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TRENDS IN U.S. WAGE INEQUALITY: RE-ASSESSING THE REVISIONISTS

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Trends in U.S. Wage Inequality: Re-Assessing the Revisionists

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A large literature documents a substantial rise in U.S. wage inequality and educational wage differentials over the past several decades and finds that these trends can be primarily accounted for by shifts in the supply of and demand for skills reinforced by the erosion of labor market institutions affecting the wages of low- and middle-wage workers. Drawing on an additional decade of data, a number of recent contributions reject this consensus to conclude that (1) the rise in wage inequality was an "episodic" event of the first-half of the 1980s rather than a "secular" phenomenon, (2) this rise was largely caused by a falling minimum wage rather than by supply and demand factors; and (3) rising residual wage inequality since the mid-1980s is explained by confounding effects of labor force composition rather than true increases in inequality within detailed demographic groups. We reexamine these claims using detailed data from the Current Population Survey and find only limited support. Although the growth of overall inequality in the U.S. slowed in the 1990s, upper tail inequality rose almost as rapidly during the 1990s as during the 1980s. A decomposition applied to the CPS data reveals large and persistent rise in within-group earnings inequality over the past several decades, controlling for changes in labor force composition. While changes in the minimum wage can potentially account for much of the movement in lower tail earnings inequality, strong time series correlations of the evolution of the real minimum wage and upper tail wage inequality raise questions concerning the causal interpretation of such relationships. We also find that changes in the college/high school wage premium appear to be well captured by standard models emphasizing rapid secular growth in the relative demand for skills and fluctuations in the rate of growth of the relative supply of college workers – though these models do not accurately predict the slowdown in the growth of the college/high-school gap during the 1990s. We conclude that these patterns are not adequately explained by either a 'unicausal' skill-biased technical change explanation or a revisionist hypothesis focused primarily on minimum wages and mechanical labor force compositional effects. We speculate that these puzzles can be partially reconciled by a modified version of the skill-biased technical change hypothesis that generates a polarization of skill demands.

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I. Introduction

A large literature documents a substantial widening of the U.S. wage structure during the 1980s (Bound and Johnson 1992; Katz and Murphy 1992; Levy and Murnane 1992; Murphy and Welch 1992; Juhn, Murphy and Pierce 1993). Wage differentials by education, by occupation and by age and experience group all rose substantially.[1] Residual wage inequality – that is, wage dispersion within demographic and skill groups – increased simultaneously. The growth of wage inequality was reinforced by changes in non-wage compensation leading to a large increase in total compensation inequality (Hamermesh, 1999; Pierce 2001). These wage structure changes translated into a pronounced rise in both household income inequality and consumption inequality, implying a marked increase in the disparities of economic well-being for U.S. families (Cutler and Katz 1991, 1992; Attanasio and Davis 1996; Karoly and Burtless 1995).

The literature documenting and interpreting these wage structure changes of the 1980s reaches two broad conclusions. First, much of the rise in U.S. earnings inequality during the 1980s appears explained by shifts in the supply of and demand for skills combined with the erosion of labor market institutions – including labor unions and the minimum wage – that protected the earnings of low and middle wage workers.[2] Second, a number of influential studies argue that the surge of inequality evident in the 1980s reflected an ongoing, secular rise in the demand for skill that commenced decades earlier and perhaps accelerated during the 1980s with the onset of the computer revolution. When this secular demand shift met with an abrupt slowdown in the growth of the relative supply of college equivalent workers during the 1980s – itself a consequence of slowing educational attainment for cohorts born after 1949 – wage differentials expanded rapidly (Katz and Murphy 1992; Autor, Katz and Krueger 1998; Goldin and Katz 2001; Acemoglu 2002).

Drawing on a more recent decade of data, however, a series of recent studies challenges the conclusions of this literature. Most notably, Card and DiNardo (2002) stake two broad claims that strongly dissent from

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[1] A substantial narrowing of gender wage differentials both overall and for all age and education groups is the primary exception to the broad pattern of a widening U.S. wage structure since 1980.

[2] See Katz and Autor (1999), Goldin and Katz (2001), and Acemoglu (2002) for overviews of this literature. See Berman, Bound and Machin (1998) and Machin and Van Reenan (1998) for international comparisons.

the view summarized above. First, they argue that the rise of inequality during the 1980s is largely explained by factors other than supply and demand, most prominently, the declining real value of the minimum wage, a view that was earlier articulated by Lee (1999).[3] Second, they conclude that the growth of U.S. earnings inequality was primarily a one-time ("episodic") event of the early 1980s, which plateaued by the mid 1980s and did not recur. Building on this line of argument, Lemieux (2005) reanalyzes the sources of residual inequality growth during the 1970s through the 1990s and draws conclusions that closely parallel Card and DiNardo (2002) for overall inequality. Lemieux argues that the marked rise of residual inequality in the 1980s was also an episodic event that is primarily accounted for by non-market factors, specifically the declining value of the minimum wage in the 1980s and the mechanical effects of the changing composition of the U.S. labor force during the 1990s (rising education and experience).

This recent ('revisionist') literature has the potential to substantially amend both the factual description of U.S. earnings inequality and its interpretation. If, as these studies suggest, the rise of U.S. earnings inequality was in large part a brief, non-recurring episode of the early 1980s, the probable causes are likely to be one-time precipitating events of that time period – most prominently, the falling minimum wage. Alternatively, if the growth of earnings inequality reflects a long-term movement towards greater dispersion of earnings, then it is more likely to be explained by fundamental, secular factors, such as the supply of and demand for skills. Thus, the traditional versus revisionist characterizations of the facts strongly motivate alternative explanations – e.g., idiosyncratic U.S. policy events versus market forces.[4]

In this paper, we reevaluate the traditional and revisionist explanations for the patterns, causes, and consequences of changes in the U.S. wage inequality over the last four decades, paying particular attention to the two main claims of the revisionist literature: (1) that the growth of inequality was an episodic rather

[3] DiNardo, Fortin and Lemieux (1996) also conclude that labor market institutions are the most important factor explaining rising wage inequality in the 1980s, but they do not attribute the majority of the increase to this factor.

[4] These explanations are not intrinsically at odds, and numerous studies focused on the experience of the 1980s support the view that institutions and market forces both contributed to rising inequality. Most work espousing this hybrid viewpoint was written without the perspective of the wage structure changes of the last 8 to 10 years, however. In light of the experience of the 1990s, Card and DiNardo (2002, p. 774) write "From the vantage point of the late 1980s, there were many reasons to find SBTC [Skill Biased Technological Change] a plausible explanation for the rise in inequality over the previous decade... Viewed from 2002, it now appears that the rise in wage inequality was an episodic event."

than secular phenomenon; and (2) that it is explained largely by non-market forces, i.e., the minimum wage and, for residual inequality, labor force composition. We use wage data from the March Current Population Surveys (CPS) covering 1963 to 2002 and from the May CPS samples for 1973 to 1978 combined with the CPS Outgoing Rotation Group (ORG) files for 1979 to 2003.

In partial support of the revisionist literature, we find that past is not prologue: the growth of wage inequality in the 1990s was considerably slower than in the 1980s, and the secular demand increases favoring more educated workers were, by our estimates, less rapid in the 1990s than in either the 1980s or 1970s. In addition, we concur with the view that the falling minimum wage was likely an important contributor to rising earnings inequality in the early 1980s, particularly for the expansion of inequality in the lower half of the earnings distribution (the 50-10 wage gap).

By contrast, we find no support for the either of the two major revisionist claims articulated above. On the first point, the growth of wage inequality is not accurately described as an episodic event. Inequality in the upper half of the male wage distribution (the 90-50 wage gap) grew rapidly and nearly-continuously from 1980 to 2003 at the rate of about 1 log points per year – a marked, secular phenomenon.[5] On the second point, the persistent rise in upper-tail inequality belies the claim that minimum wages (or other institutions protecting low wage workers) can provide a coherent explanation for the bulk of the rise in earnings inequality. This explanation is implausible both because the minimum wage appears quite unlikely to produce rising earnings dispersion above the median, and because the timing of the explanation is incorrect: upper-tail inequality rose steadily from 1980 to 2003 even though the minimum wage's freefall was largely halted after1989. In fact, the only time period during which the minimum wage appears particularly relevant to rising inequality is during 1979 to 1987.

Our rejection of the revisionist claims for overall inequality holds with equal force for residual inequality. In contrast to Lemieux (2005), we find that the growth of residual inequality is not well explained by changes in the minimum wage combined with (mechanical) labor force composition shifts.

[5] Over three-quarters of the rise in 90-10 earnings inequality from 1979 to 2003 is accounted for by the rise in the 90-50 wage gap using hourly wages for all male wage and salary workers in both the CPS ORG and March files.

Rather, we document a pronounced, secular rise in residual earnings inequality that, paralleling the rise in overall inequality, is concentrated above the median of the (residual) earnings distribution. Drawing on a more detailed analysis in Autor, Katz and Kearney (2005), we show that this rise primarily reflects changes in labor market prices, holding composition-constant. Logically, since the growth in residual inequality takes place in the upper-tail of the earnings distribution, it is also quite unlikely to be explained by a falling minimum wage. Consistent with Lemieux (2005), we confirm that changes in labor force composition exerted some upward force on residual wage dispersion during the 1990s. But this compositional effect was entirely concentrated in the lower-tail of the earnings distribution and, moreover, served primarily to offset a rapid compression of lower-tail prices. In net, we find that price changes are the dominant force in the evolution of residual wage inequality.

The remainder of this paper is organized as follows. Section II documents the evolution of the U.S. wage structure from 1963 to 2003. Section III presents some basic time series models to assess the role of demand, supply, and institutional factors in the evolution of U.S. educational wage differentials and overall wage inequality. We also explore the timing of correspondence of changes in different part of the wage distribution to changes in the federal real minimum wage. Section IV draws upon Autor, Katz, and Kearney (2005) in applying a quantile regression methodology proposed by Machado and Mata (2005) to analyze the role of labor market prices and labor force composition in changes in residual inequality. Section V recapitulates our main findings and concludes by offering some speculative alternative explanations for the observed changes in the U.S. wage structure in light of the insights gleaned from another decade of data.

II. U.S. wage structure changes over the past four decades: Key facts

To summarize the basic changes in the U.S. wage structure over the last four decades using, we draw on two, large, representative data sources collected at annual frequency: the March CPS and the combined CPS May and Outgoing Rotation Group. We describe these sources briefly here and provide many additional details in the Data Appendix. The March CPS data have been widely used in studies of U.S. wage inequality and provide reasonably comparable data on prior year's annual earnings, weeks worked, and hours worked per week for four decades. We use the March files from 1964 to 2004 (covering earnings

from 1963 to 2003) to form a sample of real weekly earnings of full-time full-year workers (FTFY), defined as those working at least 35-plus hours per week and 40-plus weeks per year. Our core sample consists of those aged 16 to 64 years in the earnings year.[6] Starting in 1976 (earnings year 1975), the March survey began collecting information on hours worked in the prior year, and this allows us to create a second March sample of hourly wage data for all wage and salary workers employed in the prior calendar year for earnings years 1975 to 2003.

We complement the March series with May CPS samples for 1973 through 1978 and CPS Outgoing Rotation Group samples for 1979 through 2003 (CPS May/ORG). We use these data to construct hourly and full-time weekly wage data for all wage and salary workers employed during the CPS sample survey reference week (limiting the weekly wage measure to the full-time subsample). Unlike the retrospective annual earnings data in the March CPS, the May/ORG data provide point-in-time measures of usual hourly or weekly earnings. We weight both March and May/ORG data by hours worked to provide a measure of the entire distribution of hours paid.[7]

As detailed in Autor, Katz and Kearney (2005) and Lemieux (2005), both March and May/ORG CPS surveys have limitations that reduce their consistency over the forty year period studied. The March CPS data are not ideal for analyzing the hourly wage distribution since they lack a point-in-time wage measure and thereby hourly wages must be computed by dividing annual earnings by the product of weeks worked last year and usual weekly hours last year. Estimates of hours worked last year from the March CPS appear to be quite noisy and data on usual weekly hours last year are not available prior to the 1976 March CPS. The May/ORG samples provide more accurate measures of the hourly wage distribution but cover a shorter time period than the March CPS. Both the March and May/ORG CPS samples have undergone various changes in processing procedures over several decades, especially involving the top-coding of high earnings

[6] We also drop from the sample (full-time) workers with weekly earnings below ½ the value of the real minimum wage in 1982 ($67 a week in 1982 dollars or $112 a week in 2000 dollars).

[7] The March data are weighted by the product of weeks worked and hours per week in the prior year. The May/ORG data are weighted by hours worked during the survey reference week. Weighting by weeks is implicit in the May/ORG sample since the probability that an individual is observed working during the sample reference week is proportional to weeks in the labor force.

and the flagging of earning imputations and algorithms used for allocating earnings to those individuals who do not answer earnings questions in the survey. These create challenges in producing consistent data series over time, which we have tried to account for to the extent possible to make the wage series comparable and time consistent. As detailed in Autor, Katz and Kearney (2005), the major redesign of the earnings questions in the CPS ORG in 1994 is likely to have created additional comparability problems that we are unable to redress.[8]

A. Trends in overall inequality

We begin laying out basic wage structure facts in Figure 1, which uses data on FTFY workers from the March CPS to illustrate the substantial overall widening of U.S. wage inequality for both men and women over the past 40 years. This figure plots the change in log real weekly wages by percentile for men and for women from 1963 to 2003.[9] The figure displays a sizable expansion of both the male and female wage distributions with the 90th percentile earners rising by approximately 45 log points (more than 50 percent) relative to 10th percentile earners for both men and women. The figure also indicates a monotonic (and almost linear) spreading out of the entire wage distribution for women and for the wage distribution above around the 30th percentile for men. Notably, women have substantially gained on men throughout the wage distribution over last four decades.[10]

To illustrate the evolution of U.S. wage inequality, we employ four inequality metrics: changes in

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[8] Both AKK 2005 and Lemieux (2005) find substantial discrepancies in trends in residual inequality measured in the May/ORG versus March samples beginning in 1994. These series closely parallel each other from 1979 to 1994, and then diverge sharply thereafter, with the March data showing a continued rise in residual inequality from 1994 to 2003 for hourly workers while the ORG data show a marked flattening. Lemieux attributes the bulk of this divergence to a differential rise in measurement error in the March sample. AKK 2005 call attention to another likely source of the discrepancy: the substantial redesign of the CPS ORG survey in 1994. This redesign substantially changed the format (and increased the complexity) of the earnings component of the survey, and was followed by a striking increase in earnings non-response: from 15.3 percent in 1993 (immediately prior to the redesign) to 23.3 percent in the last quarter of 1995 (the first quarter in which allocation flags are available in the redesigned survey), reaching 31 percent by 2001 (Hirsch and Schumacher 2004). The contemporaneous rise in the earnings imputation rate in the March survey was comparatively small.

[9] The top-coding of CPS wage data makes it not very useful for measuring changes in the very top part of the wage distribution. Thus, we symmetrically trim the top and bottom parts of the distribution in Figure 1 and focus on wage changes from the 3rd to 97th percentile.

[10] Fortin and Lemeiux (2000), Welch (2000), and Mulligan and Rubinstein (2005) present interesting exploratory efforts to reconcile the rise in male wage inequality with the simultaneous closing of the gender gap.

overall wage inequality, summarized by the 90-10 log wage differential; changes in inequality in the upper and lower halves of the wage distribution, summarized by 90-50 and 50-10 log wage gaps (which we refer to as upper and lower-tail inequality); between-group wage differentials, illustrated using the college-high school wage premium; and within-group (residual) wage inequality, summarized by the 90-10, 90-50 and 50-10 residual wage gaps conditioning on measures of education, age/experience, and gender.[11]

Figures 2a and 2b display the evolution of the 90-10 overall and residual wage gaps for males and the college-high school log wage premium for our two core samples: March FTFY 1963 to 2003 and CPS May/ORG hourly 1973 to 2003. In this figure, the estimated college-high school log wage premium represents a fixed weighted average of the college plus/high school wage gaps separately estimated for males and for females in four different experience groups. Both panels of this figure underscore a key, and oft-neglected, fact about the evolution of U.S. wage inequality over four decades, which is that the rise of inequality is not a unitary phenomenon. While all three inequality measures (aggregate, residual, and between-group) expand in tandem during the 1980s then flatten somewhat in the 1990s, these series diverged sharply in both the 1970s and the 1960s. Specifically, while overall and residual inequality were either modestly rising (March) or flat (May/ORG) during the 1970s, the college wage premium declined sharply in this decade and then rebounded even more rapidly during the 1980s. Similarly, the college wage premium expanded considerably during the 1960s, even while aggregate inequality was quiescent. These divergent patterns underscore that the 'growth of inequality' is a multi-faceted phenomenon, which is unlikely to be adequately explained by any mono-thematic explanation, be it focused on technological change or labor market institutions.

Nor does Figure 2 fully convey the complexity of these trends. Underlying the rapid growth of aggregate 90-10 inequality during the 1980s followed by a deceleration in the 1990s is a sharp divergence in inequality trends at the top and bottom of the wage distribution. This divergence is shown in Figures 3

[11] The robustness of conclusions concerning the timing of changes in overall and residual wage inequality changes to the choice of wage concept and sample is illustrated in Appendix Tables 1a and 1b which present changes over consistent sub-periods from 1975-2003 of different measures of inequality for males, females, and both combined using weekly earnings for full-time workers and hourly wages for all workers for the March CPS and May/ORG CPS.

through 5, which compare the evolution of the 90-10, 90-50 and 50-10 log hourly and full-time weekly wage gaps for males and females from 1963-2003 using the March CPS samples and for the years 1973 to 2003 using the May/ORG samples. As show in Figures 4 and 5, upper tail and lower tail wage inequality both expanded rapidly in the first half of the 1980s for both men and women in all four samples (March vs. May/ORG x FTFY vs. hourly). But the 50-10 wage gap for the most part stopped growing after 1987 – and the male hourly wage series from the CPS May/ORG shows an actual decline in the 50-10 since the late 1980s. By contrast, the 90-50 wage gap continues to grow smoothly and continuously from 1979 to the present in all wage series and for both genders. Thus, the deceleration of aggregate inequality growth since 1987 actually reflects an abrupt halt or reversal in lower-tail inequality expansion paired with a continuing secular rise in inequality above the median of the wage distribution.[12]

The divergent growth of upper and lower tail wage inequality in the 1980s and 1990s is also corroborated by microdata on wages and total compensation from the establishment-based Employment Cost Index (ECI) sample. Pierce (2001), in an analysis of the ECI microdata, finds a large rise in 50-10 and 90-50 wage (and total compensation) differentials from 1982 to 1986 for a sample combining men and women. Similar to the CPS data we use here, the rise in lower half inequality in the ECI ceases and somewhat reverses itself from 1986 to 1996, while the growth in upper-tail compensation inequality continues steadily over the next decade. As carefully documented by Piketty and Saez (2003) using Internal Revenue Service tax data, the rapid and steady growth of upper tier earnings is also strongly evidence by the rising shares of wages paid to the top 10 and top 1 percent of U.S. earners since the mid-1970s.

To summarize, the sharp growth in wage dispersion in the lower-half of the wage distribution during the early to mid 1980s seems to have been an episodic event that has not re-occurred over the past fifteen years (though it also has not been reversed). By contrast, the steady growth of wage dispersion in the upper

[12] The differences in the growth of the 90-50 and 50-10 wage differentials have previously been emphasized by Mishel, Bernstein and Boushey (2002) and are also noted by Lemieux (2005). Using decennial Census earnings data, Angrist, Chernozhukov and Fernández-Val (forthcoming) document a sharp rise in residual inequality from 1980 to 1990, with a continuing increase from 1990 to 2000 concentrated in the upper half of the wage distribution.

half of the wage distribution appears to represent a secular trend that has been ongoing for 25 years. We will have more to say about the possible sources of these diverging trends in upper and lower tail inequality in Section V. Before interpreting these relationships, we first highlight the principal trends in between-group inequality over this time period.

B. Trends in wage levels and between group inequality

Table 1 summarizes the major between-group wage structure changes by presenting mean log real wage changes by sub-period from 1963 to 2003 for various groups defined by sex, education, and potential experience. Mean (predicted) log real weekly wages were computed in each year for 40 detailed sex-education-experience groups and mean wage for broader groups are fixed-weighted averages of the relevant sub-group means, using the average share of total hours worked for each group over 1963 to 2003 as weights to adjust for compositional changes within each group.[13]

The first row indicates that composition-adjusted real wages increased by 24 log points over the full period. Wage growth was rapid in the 1960s, stagnant or declining from 1971 to 1995, and rapid in the late 1990s. The next two rows show that women gained substantially on males – by 17.9 log points over the full sample – with the growth in the relative earnings of women most concentrated in the 1979 to 1995 period.

The following six rows summarize real wage changes by educational groups. These figures highlight the expansion of educational wage differentials, with particularly large increases in the relative earnings of college graduates. The sharp differences across decades seen in Figure 2 are evident in these detailed figures, with educational wage differentials rising in the 1960s, narrowing in the 1970s, increasing sharply in the 1980s, and growing at a slightly less torrid pace in the 1990s. The bottom part of the table contrasts changes in real wages for younger and older male high school and college graduates. Experience differentials expanded for college and high school graduates with the rise for college graduates concentrated in the 1960s and 1970s and the rise for high school graduates concentrated in the 1980s.

The data in Table 1 indicate that the spreading of the wage differentials between demographic groups

[13] The March and May/ORG samples are generally considered equally valid for measuring between-group wage trends, and the March has the advantage of covering an additional decade of data.

has been less continuous – and has undergone more reversals – than has the general towards increasing aggregate wage inequality over the last four decades. These sharp transitions in between-group inequality actually make the variation easier to explain with familiar models.

III. The sources of rising inequality: Proximate causes

We now present a simple analysis of the leading proximate causes of aggregate and between-group wage inequality, focusing on supply and demand factors, unemployment and the minimum wage. We present simple time-series models of the U.S. college wage premium covering 1963 to 2002 and augment the specification to allow for an impact of changes in a key labor market institutional factor, the federal minimum wage.[14] We postpone to Section IV a discussion of factors behind the rise in residual inequality.

A. Sources of the rising college/high-school wage premium

Our illustrative conceptual framework starts with a CES production function for aggregate output Q with two factors, college equivalents (c) and high school equivalents (h):

$$(1) \qquad Q_t = [\alpha_t (a_t N_{ct})^{\rho} + (1-\alpha_t)(b_t N_{ht})^{\rho}]^{1/\rho}$$

where N_{ct} and N_{ht} are the quantities employed of college equivalents (skilled labor) and high-school equivalents (unskilled labor) in period t, a_t and b_t represent skilled and unskilled labor augmenting technological change, α_t is a time-varying technology parameter that can be interpreted as indexing the share of work activities allocated to skilled labor, and ρ is a time invariant production parameter. Skill-neutral technological improvements raise a_t and b_t by the same proportion. Skill-biased technological changes involve increases in a_t / b_t or α_t. The aggregate elasticity of substitution between college and high-school equivalents is given by $\sigma = 1/(1-\rho)$.

Under the assumption that college and high-school equivalents are paid their marginal products, we can use equation (1) to solve for the ratio of marginal products of the two labor types yielding a relationship

[14] Many studies, since Freeman (1975) and Tinbergen (1975), have used simple formal supply and demand frameworks to analyze changes in educational wage differentials. The present analysis extends earlier work in Katz and Murphy (1992) and Katz and Autor (1999), drawing on additional years of data.

between relative wages in year t, w_{ct}/w_{ht}, and relative supplies in year t, N_{ct}/N_{ht} given by

$$(2) \qquad \ln(w_{ct}/w_{ht}) = \ln[\alpha_t/(1-\alpha_t)] + \rho \ln(a_t/b_t)] - (1/\sigma)\ln(N_{ct}/N_{ht}),$$

which can be rewritten as

$$(3) \qquad \ln(w_{ct}/w_{ht}) = (1/\sigma)[D_t - \ln(N_{ct}/N_{ht})],$$

where D_t indexes relative demand shifts favoring college equivalents and is measured in log quantity units. The impact of changes in relative skill supplies on relative wages depends inversely on the magnitude of aggregate elasticity of substitution between the two skill groups. The greater is σ, the smaller the impact of shifts in relative supplies on relative wages and the greater must be fluctuations in demand shifts (D_t) to explain any given time series of relative wages for a given time series of relative quantities. Changes in D_t can arise from (disembodied) skill-biased technological change, non-neutral changes in the relative prices or quantities of non-labor inputs, and shifts in product demand.

Following the approach of Katz and Murphy (1992), we directly estimate a version of equation (3) to explain the evolution from 1963 to 2003 of the overall log college/high school wage differential series for FTFY workers from the March CPS shown in Panel A of Figure 2. We substitute for the unobserved demand shifts D_t with simple time trends and a measure of labor market cyclical conditions, the unemployment rate of males aged 25-54 years. We also include an index of the log relative supply of college/high school equivalents.[15] Our full model includes the log real minimum wage as a control variable:

$$(4) \qquad \ln(w_{ct}/w_{ht}) = \gamma_0 + \gamma_1 t + \gamma_2 \ln(N_{ct}/N_{ht}) + \gamma_3(\text{RealMinWage}_t) + \gamma_4 \text{Unemp}_t + \varepsilon_t,$$

where γ_2 provides an estimate of $1/\sigma$.

The large increase in the college wage premium over the last 40 years coincided with a substantial secular rise in the relative supply of college workers. As illustrated in Figure 6, the college graduate share of the full-time equivalent workforce increasing from about 10.6 percent in 1960 to over 31 percent in 2003. Given this rapid growth in college graduate supply, a market-clearing model will require (even more) rapid

[15] We use a standard measure of college/non-college relative supply calculated in "efficiency units" to adjust for changes in labor force composition by gender and experience groups. Full details are provided in the Data Appendix.

growth in relative demand for college workers to reconcile increasing college supply with a rising college wage premium (Table 1).

Figure 7 plots relative the series for relative supply growth over 1963 to 2003 deviated from a linear time trend. This figure reveals a clear acceleration of the relative supply of college workers in the 1970s related to the 1960s, followed by a dramatic slowdown in starting in 1982. These fluctuations in the growth rate of relative supply, paired with a constant trend growth in relative college demand, do a surprisingly effective job of explaining the modest increases in the college premium in the 1960s, the decline in the 1970s, and the sharp increase over the past two decades. The top panel of Figure 7 illustrates the explanatory power of such an approach by showing that the deviations in relative supply growth from a linear trend roughly fit the broad changes in the de-trended college wage premium from 1963 to 2002.

Table 2 presents representative regression models for the overall college/high school log wage gap following this simple approach. The first column uses the basic specification of Katz and Murphy (1992) for the 1963 to 1987 period (the period analyzed by Katz-Murphy) with only a linear time trend and the relative supply measure included as explanatory variables. Although our data processing methods differ somewhat from those of Katz and Murphy, we uncover quite similar results with an estimate of $\gamma_2 = 0.64$ (implying $\sigma = 1.56$) and with estimated trend growth in the college wage premium of 2.7 percent per annum. The lower panel of Figure 7 uses this replication of the basic Katz-Murphy model from col. (1) of Table 2 to predict the evolution of the college wage premium for the full sample period of 1963 to 2003 and compares the predicted and actual college wage gap measures.

The Katz-Murphy model does a good job of forecasting the growth of the college wage premium through 1992 (with the exception of the late 1970s) but the continued slow growth of relative supply after 1992 leads it to over-predict the growth in the college wage premium over the last decade. The most straightforward implication of this pattern is that there has been a slowdown in trend relative demand growth for college workers since 1992, as illustrated by a comparison of the models in columns (2) and (3)

of Table 2 without and with allowing for a trend break in 1992.[16] The model in column (3) covering the full 1963-2003 period indicates a significant slowdown of demand growth after 1992 but still indicates a large impact of relative supply growth with an estimated aggregate elasticity of substitution of 1.63 (1/0.612).[17]

The implied slowdown in trend demand growth in the 1990s is potentially inconsistent with a naïve SBTC story looking at the growth of computer investments since these continued most rapidly in the 1990s. But strong cyclical labor market conditions with low unemployment in expansion of the 1990s might account for some of this pattern and the differential impacts of labor market institutions such as the minimum wage might also play a role in the evolution of the college wage premium. As discussed by DiNardo, Fortin and Lemieux (1996) and Lee (1999), the real value of the U.S. minimum wage experienced a sharp decline in the 1980s and more modest movements in the 1960s, 1970s, and the past decade.

The roles of cyclical conditions and the minimum wage are examined in the augmented models illustrated in columns (4) and (5) of Table 2. The real minimum wage and prime age male unemployment rates have modest additional explanatory power in the expected directions and reduce the extent of unobserved slowdown in trend demand growth over the last decade. But the inclusion of these variables does not much alter the central role for relative supply growth fluctuations and trend demand growth in explaining the evolution of the college wage premium over the past four decades.[18] A model without the relative supply variable in column (6) leads to larger impacts of the real minimum wage but it also has much less explanatory power and generates a puzzling negative impact of prime age male unemployment on the college wage premium.

B. The college/high-school gap by experience group

As shown in Table 1, changes in the college/high school wage gap differed substantially by

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[16] This point is also noted by Autor, Katz and Krueger (1998), Katz and Autor (1999) and Card and DiNardo (2002).

[17] Similar conclusions of a significant slowdown in trend relative demand growth for college workers arise in models allowing trends breaks in any year from 1989 to 1994.

[18] The predicted effect of the minimum wage on the college/high-school gap in these regressions is not economically large. The real minimum wage fell by 34.7 log points between 1979 and 1989, which implies an increase the college/high school gap of 3.9 log points using the point estimate in column 5 of Table 2. In actuality, the college/high-school gap rose by 13.7 log points during 1979 to 1989.

age/experience groups over recent decades, with the rise in the college/high-school gap concentrated among less experienced workers in the 1980s (see the bottom rows of Table 1). We illustrate this pattern in Figure 8 through a comparison of the evolution of the college premium (panel A) and college relative supply (panel B) for younger workers (those with 0-9 years of potential experience) and older workers (those with 20-29 years of potential experience). As is visible in this figure, the return to college for younger workers has increased much more substantially since 1980 than for older workers. To the extent that workers with similar education but different ages or experience levels are imperfect substitutes in production, one should expect age-group or cohort-specific relative skill supplies – as well as aggregate relative skill supplies – to affect the evolution of the college-high school by age or experience as emphasized in a careful analysis by Card and Lemieux (2001). Consistent with this view, the lower panel of Figure 8 shows a much more rapid deceleration in relative college supply among younger than older workers in the mid to late 1970s.

In Table 3, we take fuller account of these differing trends by estimating regression models for the college wage by experience group that extend the basic specification in equation (4) to include own experience group relative skill supplies.[19] The first two columns of Table 3 present regressions pooled across 4 potential experience groups (those with 0-9, 10-19, 20-29, and 30-39 years of experience) allowing for group-specific intercepts but constraining the other coefficients to be the same for all experience groups. These models estimate:

$$(5) \qquad \ln(w_{cjt}/w_{hjt}) = \beta_0 + \beta_1[\ln(N_{cjt}/N_{hjt}) - \ln(N_{ct}/N_{ht})] + \beta_2 \ln(N_{ct}/N_{ht}) + X_t\beta_3 + \delta_j + \eta_{jt},$$

where j indexes experience group, the δ_j are experience group main effects, and x_t includes measures of time trends and other demand shifters. This specification arises from an aggregate CES production function in college and high school equivalents of the form of equation (1) where these aggregate inputs are themselves CES sub-aggregates of college and high school labor by experience group (Card and Lemieux 2001). Under these assumptions, $-1/\beta_2$ provides an estimate of σ the aggregate elasticity of substitution

[19] An analogous set of regressions to explain the evolution of the college wage premium by age groups is presented in Appendix Table 2.

and $-1/\beta_1$ provides an estimate of σ_E, the partial elasticity of substitution between different experience groups within the same education group.

The estimates in the first two columns of Table 3 indicate substantial effects of both own-group and aggregate supplies on the evolution the college wage premium by experience group. While the implied estimates of the aggregate elasticity of substitution in the Table 3 models are very similar to the aggregate models in Table 2, the implied value of the partial elasticity of substitution between experience groups is around 3.42 (somewhat lower than the estimates in Card and Lemieux 2001). These estimates indicate that differences in own-group relative college supply growth go a substantial distance towards explaining variation across experience groups in the evolution of the college wage premium in recent decades. For example, as seen in Figure 8, from 1980 to 2002 the college wage premium increased by 28.8 log points for the 0-9 year experience group and by 19.5 log points for the 20-29 year experience group. Over the same period the own group relative college supply for the 0-9 year experience group grew by 25.1 log points less rapidly than for the 20-29 year experience group. Thus, using the implied own-group relative inverse substitution elasticity of -0.292 in column (1) of Table 3, we find that the slower relative supply growth for the younger (0-9 year) experience group explains most (79% or 7.3 log points of a 9.3 log point gap) of the larger increase in the college premium for the younger than for the older (20-29 year) experience group.

The final four columns of Table 3 present analogous regression models of the college wage premium separately estimated by experience group. Trend demand changes and relative skill supplies play a large role in changes in educational differentials for younger and prime age workers. The post-1992 slowdown in trend demand growth is apparent for the youngest experience group but not for prime age workers. The college wage premium for younger workers appears more sensitive to own group and aggregate relative skill supplies than the premium for older workers. We also find that the real minimum wage is a significant determinant of changes in the college wage premium for younger workers (those with less than 20 years of potential experience), but, plausibly, does not appear important for more experienced workers groups.

In summary, these estimates suggest that a simple demand and supply framework that is predicated on a

secular increase in relative demand for educated workers can account for many of the key patterns in between-group inequality, including the contraction and expansion of the college-high school gap during the 1970s and 1980s and the differential rise in the college/high-school gap by experience group in the 1980s and 1990s. However, the model over-predicts the rise in between group inequality after 1992, a puzzle to which we return to below.

C. The role of the minimum wage

A number of prominent studies, including Lee (1999), Card and DiNardo (2002) and Lemieux (2005), have emphasized the central role of the minimum wage in explaining the rise of U.S. earnings inequality. Yet, our simple models above do not suggest that the minimum wage (or fluctuations in aggregate unemployment) have played an important role in the evolution of educational wage differentials in recent decades, except in the case of inexperienced workers. Why do our conclusions differ? The discrepancy arises from the disjuncture between trends in between-group inequality (as measured by the college/high-school gap) and trends in aggregate and residual inequality. As depicted in Figure 2, overall inequality was flat during the 1970 while between-group inequality fell; conversely, as between group-inequality rose during the 1990s, aggregate inequality stabilized (at least in some data sources). In fact, it is only during 1979 to 1987 that between-group and aggregate inequality move closely together. It is therefore unlikely that any simple unicausal framework will be successful at explaining all trends simultaneously. As shown below, while supply-demand models are reasonably successful at explaining movements in between-group inequality, they are far less effective at explaining other inequality metrics.

In the same spirit as our models for the college/high-school earnings gap above, we provide in this section a simple time-series analysis for the proximate sources of the growth of aggregate hourly wage inequality, as measured by the 90/10, 90/50, and 50/10 log wage ratios. As emphasized by Card and DiNardo (2002), there is a striking time series relationship between the real value of the federal minimum wage and hourly wage inequality, as measured by the 90-10 log earnings ratio. This relationship is depicted in Figure 9. A simple regression of the 90-10 log hourly wage gap from the May/ORG CPS for the years 1973 to 2003 on the real minimum wage (deflated by the PCE deflator) and a constant yields a coefficient

of -0.71 and an R-squared of 0.69. Based in part on this tight correspondence, Card and DiNardo (2002) and Lemieux (2005) argue that much of the rise in overall and residual inequality over the last two decades may be attributed to the minimum wage. Using a cross state analysis of minimum wage levels and earnings inequality, Lee (1999) also concludes that were it not for the falling U.S. minimum wage, there would have been no rise in inequality during the 1980s.

A potential problem for this line of argument is that the majority of the rise in earnings inequality over the last two decades occurred in the upper half of the earnings distribution (see Figures 3 through 5 and Appendix Table 1a). Since it is not obvious why a declining minimum wage would cause upper-tail earnings inequality to rise, this observation suggests that the minimum wage is unlikely to provide a complete explanation for overall inequality growth.[20] A further concern about the validity of this explanation is raised by comparing minimum wage levels with upper and lower-tail inequality. As shown in the upper panel of Figure 10, the level of the minimum wage is highly correlated with lower-tail earnings inequality between 1973 and 2003; a 1 log point rise in the minimum is associated with 0.27 log point compression in lower tail inequality. Somewhat surprisingly, the minimum wage is also highly correlated with upper tail inequality. Over this time interval, a 1 log point rise in the minimum is associated with a 0.44 log point compression in upper tail inequality (Figure 10, lower panel).

These bivariate relationships may potentially mask other confounds. To explore the relationships in slightly greater detail, we estimate in Table 4 a set of descriptive regressions for 90-10, 90-50 and 50-10 hourly earnings inequality over 1973 to 2003. In addition to the minimum wage measure used in Figures 9 and 10, these models add a linear time trend, a measure of college/high-school relative supply (calculated from the May/ORG CPS), the male prime-age unemployment rate (as a measure of labor market tightness), and in some specifications a post-1992 time trend, reflecting the estimated trend reduction in skill demand in the 1990s. The main finding from these models, visible in Table 4, is that the relationship between the minimum wage and both upper and lower-tail inequality is robust, although magnitudes are substantially

[20] Moreover, as shown in the upper panel of Figure 9, the slide in the real minimum wage was halted after 1989 and partly reversed over 1989 to 1998, yet the trend rise in upper-tail inequality continued unabated.

reduced by the inclusion of other covariates (particularly the time trend for upper-tail inequality). In a specification that includes a linear time trend, the college/high school supply measure, and the prime-age unemployment rate variable, the minimum wage measure has a coefficient of -0.26 for lower tail inequality and a coefficient of -0.08 for upper tail inequality (both significant). Aside from the time trend and the minimum wage measure, the other explanatory variables in these regressions are insignificant.[21]

These patterns suggest that the striking time series correlation between minimum wages and inequality is unlikely to provide an accurate account of the causal effect of the minimum wage on earnings inequality. Indeed, we view the relationship between the minimum wage and upper tail inequality as potential evidence of spurious causation. While we concur with Card and DiNardo (2002) that the decline in the real minimum wage during the 1980s likely contributed to the expansion of lower tail inequality – particularly for women – the robust correlation of the minimum wage with upper tail inequality suggests that other factors are at work.[22] One possible explanation for this link is that federal minimum wage shifts during these decades were in part a response to current economic conditions. The rapid fall in the minimum wage during 1980s took place during a deep recession. The legislated increases in the minimum wage in 1990 to 1991 and 1996 to 1997 occurred during relatively better economic conditions. If these macroeconomic shocks also directly affected earnings inequality, this would in part explain the coincidence of falling minimum wages and rising upper tail inequality.[23]

IV. Rising residual inequality: The role of composition and prices

[21] When we replace aggregate with residual inequality in comparable models, we also find a significant relationship between the minimum wage and upper and lower-tail inequality, though the estimates for lower-tail inequality are substantially more precise.

[22] Indeed, Lee (1999) also noted a puzzling relationship between the 'effective' state minimum wage (that is, the log difference between the state median and the state minimum) and state level upper-tail inequality. Opposite to the simple time-series regressions above, Lee's cross-state analysis finds that increases in the effective state minimum wage appear to reduce upper-tail inequality, both for males and for the pooled-gender distribution (see Lee, 1999, Table II). Due to this puzzling result, Lee suggested caution in causally attributing trends in male and pooled-gender earnings inequality to the minimum wage. Ironically, several studies appear to have misread Lee's findings as suggesting that 'spillovers' from the minimum wage can also account for the rise in upper-tail inequality. In fact, the Lee results imply spillovers of the opposite sign.

[23] In a similar vein, Acemoglu, Aghion and Violante (2001) argue that the decline in union penetration in the United States and the United Kingdom is partly explained by changing skill demands that reduced the viability of rent sharing bargains between high and low skill workers.

The educational attainment and labor market experience of the U.S. labor force rose substantially over the last 25 years as the large 1970s college cohorts reached mid-career during the 1990s. Tabulations from the CPS May/ORG samples indicate that the full-time equivalent employment share of male workers with a college degree rose from less than one-fifth to fully one-third of the U.S. male labor force between 1973 and 2003 and the employment share of workers with high school or lower education fell by one third (from 62 to 41 percent). The mean potential experience of workers with high school or greater education increased by 2 to 6 years between 1973 and 2003, with the largest gains experienced by the most educated groups.

As emphasized by Lemieux (2005), these shifts in labor force composition may have played a role in recent changes in measured wage inequality. The canonical Mincer (1974) earnings model implies that earnings trajectories fan out as workers gain labor market experience. Hourly wage dispersion also is typically higher for college graduates than for less-educated workers. Thus, changes in the distribution of education or experience of the labor force can lead to changes in wage dispersion. These compositional effects are distinct from the standard price effects arising from shifts in supply-demand and institutional factors. Holding market prices constant, changes in labor force composition can mechanically raise or lower residual earnings dispersion simply by altering the employment share of worker groups that have more or less dispersed earnings. Similarly, changes in workforce composition can also raise or lower overall earnings dispersion by increasing or reducing heterogeneity in observed skills (Juhn, Murphy and Pierce 1993). These observations suggest that measured earnings dispersion may change due to the mechanical impact of composition without any underlying change in market prices.[24]

Following such an approach, Lemieux (2005) finds that most of the growth in residual wage dispersion in the U.S. from 1973 to 2003 – and all of the growth after 1988 – is explained by mechanical effects of changes in workforce composition rather than shifts in residual inequality within defined skill groups (what we call price effects). Lemieux concludes that the rise in residual earnings inequality is mainly attributable to institutional factors during the 1980s – especially the falling real minimum wage – and to mechanical

labor force composition effects since the late 1980s.

In this section, we reassess these conclusions, drawing on a detailed analysis in Autor, Katz and Kearney (2005, hereafter 'AKK 2005'). Our approach is conceptually similar to Lemieux (2005), with two differences. First, and potentially less important, we apply and extend a quantile-regression-based modeling technique proposed by Machado and Mata (2005) to simulate counterfactual earnings distributions that account for the separate contributions of prices and composition. As discussed in AKK 2005, this quantile simulation is, in our view, a simple and conceptually appealing alternative to the DiNardo, Fortin and Lemieux (1996) kernel reweighting approach, but its substantive differences with Lemieux (2005) are not consequential for our conclusions.

Of greater importance, we separately analyze the contribution of prices and composition to upper and lower tail earnings inequality. It is this detailed focus on divergent trends in upper and lower-tail inequality that causes us to draw distinctly different conclusions from Lemieux (2005).

A. Implementation

Let $Q_\theta(w|x)$ for $\theta \in (0,1)$ denote the θ^{th} quantile of the distribution of the log wage (w) given the vector x of covariates. Using our May/ORG CPS samples for 1973 to 2003, we estimate quantile regression (QR) models for log hourly wages by gender and year of the form:

$$(6) \qquad Q'_\theta(w|x) = x'\beta_t(\theta),$$

where x is a $k \times 1$ vector of covariates and $\beta_t(\theta)$ is a conformable vector of quantile regression (QR) coefficients. This covariate vector includes five schooling completion categories (high school, high school graduate, some college, bachelor's degree, and post-college schooling), thirteen potential experience categories ranging (from 0 to 38 years in 3-year increments) and a complete set of interaction terms among schooling and experience categories.[25]

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[24] Labor force composition will also affect wage dispersion though general equilibrium effects of quantities on prices. We focus here on the partial equilibrium effects and provide a fuller discussion of general equilibrium effects in Autor, Katz, and Kearney (2005).

[25] We estimate this model separately for each year and gender for quantiles [0.1, 0.3,...,99.7,99.9] at intervals of one-fifth of a centile, with one additional model fit for the median (quantile 50.0).

For purposes of our analysis, the key feature of the QR model is that it can be used to partition the observed wage distribution into 'price' and 'quantity' components – that is, components attributable to the distribution of the $x's$ and components due to the (estimated) matrix of prices, $\hat{\beta}_\tau(\theta)$. This division is similar to a standard Oaxaca-Blinder procedure using OLS regression coefficients, with the key difference that the OLS model only characterizes the central tendency of the data (i.e., the conditional mean function, describing 'between-group' inequality). In contrast, the conditional quantile model characterizes both the central tendency of the data (in this case, the median) and the dispersion of the outcome variable conditional on x, i.e., the wage 'residuals.'

This latter feature is critical for estimating the impact of composition on the shape of the overall or residual wage distribution. As shown by Machado and Mata (2005), the QR regression coefficients obtained from estimates of (6) can be used to simulate the counterfactual distribution of wages that would prevail if labor force composition were given as in time period t and labor market prices were given as in time period τ. This simulation is accomplished by applying the labor force composition data $g_t(x)$ from a given time period t to the price matrix $\hat{\beta}_\tau(\theta)$ from any other time period τ to form a counterfactual wage distribution. Because the $\hat{\beta}_\tau(\theta)$ matrix describes the conditional distribution of wages for given values of x, this simulation captures the effects of composition on both between-group and residual inequality.[26]

As demonstrated in AKK 2005, the QR model can be readily extended to further decompose the price component of aggregate inequality into within- and between-group price subcomponents, an idea also developed by Melly (forthcoming). Specifically, we define the coefficient vector $\hat{\beta}(50)$ as a measure of between-group inequality, and we refer to it as $\hat{\beta}^b \equiv \hat{\beta}(50)$. $\hat{\beta}^b$ serves a role akin to $\hat{\beta}_{OLS}$ in a conventional Oaxaca-Blinder decomposition. In the conventional application, $\hat{\beta}_{OLS}$ provides a measure of between-group inequality because it estimates the central tendency of the data conditional on x. In our

[26] The details of this simulation procedure are given in AKK 2005, along with a 'proof of concept' demonstrating the efficacy of this method for accurately capturing levels and trends in overall and residual inequality.

application, the central tendency measure is the conditional median, estimated by $\hat{\beta}^b$.[27]

Following this logic, we define a measure of within-group inequality as the difference between the estimated coefficient vector $\hat{\beta}(\theta)$ and the median coefficient vector $\hat{\beta}^b$:

$$(7) \qquad \hat{\beta}^w(\theta) \equiv [\hat{\beta}(\theta) - \hat{\beta}^b] \text{ for } \theta \in (0,1).$$

By construction, $\hat{\beta}^w(50) = 0$. Hence, the residual quantile coefficient matrix is purged of 'between-group' inequality, and measures the expected dispersion of w at any given value of x, holding the conditional median at zero. By applying the coefficient matrix $\hat{\beta}^w$ to the distribution of covariates, $g(x)$, we can calculate the (estimated) dispersion of w that is exclusively attributable to residual inequality.

Putting the pieces together, we write the observed wage distribution in time period t as $f_t(\hat{w}_t) \equiv f(g_t(x), \hat{\beta}_t^b, \hat{\beta}_t^w)$, where $f(g_t(x))$ is the distribution of observable characteristics (education, potential experience and age), and $\hat{\beta}^b$ and $\hat{\beta}^w$ are vectors of 'between-group' and 'within-group' prices estimated from quantile regressions.

B. Residual inequality: The role of composition and prices

We now use the extended Machado-Mata quantile tool to implement the Lemieux (2005) decomposition for residual inequality. The results of this decomposition for the May/ORG sample at once confirm the Lemieux (2005) findings and, we believe, substantially alter their interpretation.

In our notation, the Lemieux approach (superscripted by L below) estimates the following quantity:

$$(8) \qquad \Delta Q_\theta^L = Q_\theta(f(g_\tau(x), \beta^b = 0, \beta_t^w)) - Q_\theta(f(g_t(x), \beta^b = 0, \beta_t^w)).$$

The counterfactual contribution of changing labor force composition to the change in the θ^{th} quantile of the residual distribution is estimated by replacing the x distribution from period t with the x distribution from period τ while holding residual prices at their base (period t) level and setting between-group prices to zero.

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[27] As demonstrated in AKK 2005, the difference between conditional means and medians is not empirically important in our application.

We evaluate the importance of compositional shifts for changes in residual wage inequality by applying the labor force composition data, $g_t(x)$, from each sample year to the within-group price series ($\hat{\beta}^w$) from four different years: the contemporaneous year – thus producing that year's observed level of residual inequality – and the price series for years 1973, 1988 and 2003. Residual inequality statistics (90/10, 90/50, 50/10) for these counterfactual densities for males using the May/ORG CPS samples are plotted in Figures 11a and 11b. The differences in the vertical height of each series within a given year in the figure reflect the effect of within-group prices on residual earnings inequality, holding labor force composition at the appointed year's level. The over-time change in the level of each series (moving along the x-axis) reflects the effect of changes in labor force composition, holding prices at their 1973, 1988 or 2003 level.

It bears emphasis that this counterfactual exercise depends on the maintained partial-equilibrium assumption that prices and quantities can be treated as independent. While convenient, this assumption is economically unappealing, and, moreover, is precisely opposite in spirit to our supply-demand analysis in Section III. We nevertheless view this counterfactual exercise as useful because it allows us to directly assess the substantive conclusions of Lemieux (2005), taking the modeling assumptions as given.

Figure 11a shows that holding composition constant at the 1973, 1988 or 2003 level, male 90/10 residual wage inequality rose sharply between 1973 and 1988 (compare the height of the 1973 versus 1988 series). Between 1988 and 2003, however, residual 90/10 inequality contracted by about 15 to 30 percent of its original rise, holding composition constant. This confirms the finding of Lemieux (2005) that residual inequality plateaued or contracted after 1988. But the vertical differences among these three counterfactual series at each point along the x-axis reveal that changing 'residual prices' are primarily responsible for the rise and then contraction in residual inequality in the first and second halves of the sample. These shifts in residual inequality occur with composition constant.[28]

[28] In Katz and Autor (1999), we report that residual inequality also rose in the May CPS between 1973 and 1979. As Lemieux (2005) correctly points out, this conclusion derives from a comparison of a 1973 CPS file excluding allocated earnings observations and a 1979 file including allocated observations. Once allocators are excluded from both samples, we find, consistent with DiNardo, Fortin and Lemieux (1996) and Lemieux (2005) that there is no rise in residual inequality in the May CPS between 1973 and 1979.

What role did composition play? As seen in the shallow upward slopes of each counterfactual series (moving along the x-axis), compositional shifts also contributed to rising residual inequality (holding residual prices constant), particularly after 1988. But these compositional shifts are modest relative to the price effects. These modest effects can, however, be reconciled with Lemieux's finding that compositional changes explain the full rise in residual inequality after 1988. The reconciliation is found by studying the actual net rise in estimated residual inequality from 1988 to 2003 (series labeled "observed residual"). The observed rise in residual inequality over 1988 to 2003 is quite small; we estimate it at 0.4 log points. Accordingly, the relatively modest contribution of compositional shifts to rising residual inequality can be said to explain all (in fact, more than all) of the observed rise in residual inequality for this period. But what Figure 11 reveals is that the offsetting effects of price contractions were economically as or more important than these compositional shifts.

These points apply with even greater force to upper and lower tail residual inequality treated separately. As shown in Figure 11b, composition-constant lower-tail inequality unambiguously contracted between 1988 and 2003. This is entirely a price effect; composition effects work in the opposite direction. By contrast, upper-tail residual inequality continued to rise substantially between 1988 and 2003, as it did in the prior half of the sample. In fact, only when labor force weights from 1973 are used in place of labor force weights from 1988 or 2003 does this rise appear muted for males.

As shown in AKK 2005, the reason that this latter conclusion is sensitive to weighting is that 90/10 inequality rose among better educated workers after 1988 but fell among less educated workers during the same interval. Since the labor force in 1973 was composed of considerably fewer highly educated workers and considerable more high school graduates and dropouts than in 1988 or 2003, the use of 1973 labor force characteristics puts substantial weight on the groups that experienced falling inequality and puts proportionately less weight on the groups that experienced rising inequality, suggesting a smaller rise in counterfactual residual inequality than the same comparison made using 1988 or 2003 characteristics. The sensitivity of conclusions about trends in inequality to the choice of weights cautions researchers against focusing exclusively on one set of labor force characteristics.

In summary, we find that composition plays only a secondary role in explaining the time patterns of residual inequality for males in the CPS May/ORG. We show in AKK 2005 that similar results hold for residual inequality for females and for overall wage inequality for both males and females.[29] The ongoing rise of upper-tail inequality and the rise and then stagnation of lower-tail inequality are both primarily accounted for by changing labor market prices.

V. Conclusion: Interpreting changes in the wage structure

The incorporation of data covering the labor market developments of the full 1990s and the beginning of the 21st century provides a new opportunity to assess conclusions concerning explanations for the evolution of the U.S. wage structure. Our analysis provides two clear continuities and two clear discontinuities with earlier work covering wage structure changes through the early 1990s.

The first reinforcing finding is that a simple framework emphasizing shifts in the relative demand for and relative supply of skills remains quite helpful for understanding changes in "between group" wage inequality. As emphasized by earlier work (including Katz and Murphy 1992; Murphy and Welch 1992; Autor, Katz, and Krueger 1998; and Card and Lemieux 2001), the evolution of the college-high school wage premium over the last four decades – a modest rise in the 1960s, a decline in the 1970s, and a steep rise in the 1980s continuing a more moderate rate in the 1990s – is well-explained by a strong and rather steady trend growth in the relative demand for college versus non-college labor overlaid with fluctuations in the rate of growth of the relative supply of college equivalents (particularly the surge in new college graduates of the 1970s and sharp slowdown of relative supply growth starting in the early 1980s).

Furthermore, differences in group-specific relative supply changes help explain differences by experience (or age) groups in the evolution of the college wage premium over the past couple decades. Card and Lemieux (2001) reach similar conclusions concerning the role of secular relative demand growth combined with relative supply fluctuations for explaining aggregate and age-group specific movements in the college wage premium for Canada and the United Kingdom. And Fortin (2004) finds an important role

[29] AKK 2005 also find an even large role for within-group price changes in the upper-half of the wage distribution for rising wage inequality after 1988 in analyses using March CPS rather than May/ORG hourly wage data.

for relative demand and supply shifts in explaining U.S. cross-state patterns of the evolution of the college wage premium over the last 25 years.

The second continuity with earlier findings is the almost linear rise in upper-tail wage inequality (including the 90-50 wage differential) from 1980 to the present for males and females. This persistent rise, visible in both CPS March and May/ORG wage data (as well as earnings and compensation data from the U.S. Census and Employment Cost Index), is a robust fact not readily reconciled with explanations for growing inequality that focus on minimum wages or spurious compositional effects.

The first key discontinuity with findings based on data through the end of the 1980s is that there appears to be a significant deceleration in the trend growth of relative demand for college workers starting around 1992. This pattern represents a puzzle for naïve versions of explanations focusing exclusively on skill biased technical change (SBTC). An implied acceleration in the rate of growth of the relative demand for skills in the 1980s is often attributed to the computer revolution. This hypothesis would not predict a deceleration in relative demand in the 1990s given the continued rapid spread of information technology.[30]

The second discontinuity is the near plateau (and, in some data sources, decline) in lower-tail wage inequality in the 1990s. In the 1980s wage inequality (with the notable exception of gender wage differentials) expanded substantially along many dimensions – upper tail and lower tail inequality; between-group and within-group inequality. In contrast, wage structure changes were more asymmetric in the 1990s particularly with a divergence in upper-tail and lower-tail inequality trends.

We conclude, as emphasized by Katz and Autor (1999) and DiNardo and Card (2002), that it is unlikely that a single factor (whether it be SBTC, the minimum wage, declining unionization, immigration, international trade, or shifts in labor force composition) can account for the full pattern of changes in the U.S. wage structure over the past several decades. A key reason, underscored by Figure 2, is that changes in different measures and components of wage inequality have distinct timings suggesting the possibility of their being at least partially affected by distinctive sources.

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[30] See also Machin (2002) for a discussion and interpretations of these trends.

A. Some speculative explanations

We finish by offering some brief comments on the competing explanations for changes in the U.S. wage structure. Secular demand growth for more educated workers driven by shifts in product demand and technology combined with fluctuations in relative skill supplies play a major role in the evolution of educational wage differentials. The real federal minimum wage appears to be an important factor in explaining the sharp timing of movements in lower-tail wage inequality for women, and, to a lesser degree, for men. But the minimum wage explanation fails to account for the large and persistent rise in upper tail wage inequality that has been the largest component of rising overall wage inequality since 1980. The strong time series correlation of the evolution of the real minimum wage and upper-tail wage inequality leads one to be skeptical of simple time series correlations of the real minimum wage and alternative inequality measures and is suggestive of the political endogeneity of the minimum wage.

We believe the major new puzzle introduced by the last decade's experience is the asymmetric trends in upper and lower tail inequality. We speculate that two classes of explanations may be plausible.[31] The first involves macro factors – tight labor markets in particular – that disproportionately 'raised the boats' of low-wage workers in the 1990s and offset secular labor market shifts against less-skilled workers. We doubt this can be a complete explanation, and our simple time series models above do not suggest that aggregate labor market conditions have played an important role in the evolution of between-group or overall inequality.

A second class is the one offered by Autor, Levy and Murnane 2003 ('ALM' hereafter) and amplified by Goos and Manning (2003) and Spitz (2005). Skill Biased Technical Change is probably an insufficiently nuanced name for the shifts in skill demands that we believe were induced or abetted by the rapid price declines in computer technology over the last three decades. As ALM argue, computerization is likely to have had non-monotonic impacts on the demand for skill throughout the earnings distribution: sharply raising demand for the cognitive and interpersonal skills used by educated professionals and managers; reducing demand for clerical and routine analytical skills that comprised many middle-educated white collar

[31] Piketty and Saez (2001) and Saez and Veal (2005) propose a third explanation for the evolution of (upper-tail) earnings inequality: changes in social norms. We are unclear what types of evidence would weigh for or against it.

jobs; and reducing demand for routine manual skills of many previously high-paid manufacturing production jobs.[32] Somewhat paradoxically, computerization has probably had little impact on the demand for the non-routine manual skills used in many 'low-skilled' service jobs such as health aides, security guards, orderlies, cleaners, servers, etc.[33] The ALM framework suggests that computerization (among other forces) may have raised demand for skill among higher-educated workers, depressed skill demands for 'middle-educated' workers, and left the lower echelons of the wage distribution comparatively unscathed.[34] Goos and Manning (2003) label this process a "polarization of work," and argues that it may have contributed to a hollowing out of the wage distribution in the United Kingdom during 1975 to 2000. Spitz (2005) also reports a similar polarization of job for the former West Germany during 1979 to 1999.[35]

To provide a rough assessment of the applicability of this polarization hypothesis to data for the United States, we use the Dictionary of Occupational Titles (DOT) task measures developed by ALM to examine predicted changes in employment by decile of the wage distribution over 1960 to 2000. To construct an employment demand index, we pair data from the 1960 Census of Population with the five broad ALM measures of job tasks: routine manual, routine cognitive, non-routine manual, non-routine analytic, and non-routine interactive tasks.[36] Using the paired Census-DOT sample, we calculate the mean level of task input in each decile of the wage distribution in 1960. We refer to these measures as the 'task intensity' in each wage decile, and we take them as fixed over the sample. Using Census samples for 1960, 1970 and 1980, and CPS MORG samples for 1980, 1990, and 2000, we calculate the proportionate economy-wide mean change in each measure of task input for all employed workers (weighting by the product of sampling weights and labor supply) over each decade. We multiply these aggregate proportionate changes by the

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[32] A related earlier model along these lines is developed in Juhn (1994).

[33] See also Levy and Murnane (2004) for numerous, paradigmatic examples.

[34] Welch (2000) and Weinberg (2000) argue that these technical changes are particularly likely to have been favorable to demand for female labor.

[35] Acemoglu (1999) offers an alternative theory of job polarization based on endogenous changes in production techniques as a response to a rise in the availability of skilled labor. See also related models in Acemoglu (1998) and Beaudry and Green (2003). Lewis (2005) provides evidence that manufacturing firms endogenously adopt production technologies with differing skill demands as a response to changes in local skill supplies.

[36] Task definitions are explained in detail in Autor, Levy, and Murnane (2003).

1960 task intensity in each wage decile and sum over the five task measures to estimate the aggregate predicted change in task demand by decile. Hence, if a wage decile is heavily 'tasked' in routine cognitive activities, an economy-wide decline in input of routine cognitive tasks is predicted to particularly depress task demand in that decile. Finally, to convert task changes by decile into a relative (cross-decile, within decade) measure, we express the change in each decile as a share of the total (absolute) predicted changes observed in all deciles over the decade and normalize by subtracting multiplying by 100 and subtracting 10. If, for example, predicted employment shifts were equally distributed over all 10 deciles in a decade, each would have a value of zero in our index. If all of the (relative) employment shifts were concentrated in two deciles (for example, the 1st and 10th deciles shifted by offsetting amounts), then one would have a value of 50 and other a value of negative 50.

The results of this exercise, depicted in Figure 12, indicate a notable twist in predicted employment demand by decile over four decades. During the 1960s, demand shifts are relatively uniformly distributed across deciles of the distribution, with the lowest relative growth in the highest three deciles. This pattern changes noticeably thereafter. In the 1970s, demand shifts are essentially monotonically increasing by decile. During the 1980s, positive demand shifts become even more concentrated in the top three deciles, while the most negative demand shifts are found in the bottom and middle of the distribution. In the 1990s, this twisting becomes most evident: essentially all relative demand growth in the most recent decade is concentrated in the upper three deciles, whereas relative shifts are relatively uniformly negative among the six deciles below. Notably, demand growth in the lowest decile appears less negative than in the four deciles above during this decade, consistent with modest polarization of demand.

We view these results as suggestive of a growing twist in skill demand that is at least roughly consistent with the polarization hypothesis. We stress that this simple analysis does not provide a rigorous assessment of the hypothesis but merely provides a simple illustration of its potential relevance.[37] In a more thorough analysis of changes in job quality for the 1970s and 1980s, Gittleman and Howell (1995) present evidence

[37] Note that the estimated differential 'demand' shifts across task deciles calculated from this exercise do not appear particularly large, which likely further underscores the coarseness of the technique.

that employment shifts during the 1980s were strongly biased towards upper-tier jobs and against middle-tier jobs, but had essentially no impact on employment in the bottom third of the job-quality distribution (see also Acemoglu 1999).

In conclusion, the asymmetric pattern of changes of recent changes in upper- and lower-tail wage inequality raise puzzles for both the traditional and revisionist interpretations of changes in the wage structure. Although we speculate that these trends can be partially reconciled by a reinterpretation of the skill biased technical change hypothesis along the lines developed by Autor, Levy and Murnane (2003) and Goos and Manning (2003), it is certain that there is more to be understood about the interactions among supply and demand, labor market institutions, and macroeconomic conditions than is accommodated by this simple model. Equally important, we suspect that that the factors contributing to the evolution of U.S. inequality in the decade(s) to come will differ from those identified above. While outsourcing, immigration and international trade have generally been attributed secondary roles in the recent evolution of U.S. earnings, these factors appear likely to become increasingly important, due both to rapid economic development in Asia and improvements in computer and communications technology that have dramatically reduced the costs of large scale international trade in goods and services. Devising innovative and rigorous means to evaluate the impacts of these evolving forces on inequality and economic well-being constitutes a major agenda item for research in this field.

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VII. Data appendix

A. Basic processing of May/ORG CPS data

We use the May CPS for 1973 to 1978 and the CPS Merged Outgoing Rotation Groups for years 1979 to 2003. All samples include wage/salary workers ages 16 to 64 with 0 to 39 years of potential experience in current employment. Earnings weights are used in all calculations. Full-time earnings are weighted by CPS sampling weights. Hourly earnings are weighted by the product of CPS sampling weights and hours worked in the prior week. Full-time earnings are the logarithm of reported usual weekly earnings. Hourly wages are the logarithm of reported hourly earnings for those paid by the hour and the logarithm of usual weekly earnings divided by hours worked last week (not usual weekly hours) for non-hourly workers. We use hours last week instead of usual weekly hours because usual weekly hours is not consistently available: starting with the CPS redesign in 1994, workers who report that their weekly hours vary are not asked to report usual weekly hours, yielding a non-report rate of 7.0 to 8.5 percent of workers in 1994 to 2003. To check sensitivity, we have tabulated and plotted overall and residual inequality measures using imputed usual weekly hours in place of hours last week in all years 1973 – 2003. This has little impact on our results.

Topcoded earnings observations are multiplied by 1.5. Full-time earnings of below $67/week in 1982$ ($112/week in 2000$) and hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000$) are dropped, as are hourly wages exceeding 1/35th the topcoded value of weekly earnings. All earnings numbers are deflated by the chain-weighted (implicit) price deflator for personal consumption expenditures. Allocated earnings observations are excluded in all years, except where allocation flags are unavailable (January 1994 to August 1995). As discussed by Hirsch and Shumacher (2004), only about 25 percent of allocated observations in the MORG CPS are actually flagged as allocated between 1989 and 1993. Following Lemieux (2005), we identify and drop non-flagged allocated observations by using the unedited earnings values provided in the source data.

B. Basic processing of March CPS data

We use the March Current Population Survey for earnings years 1963 to 2003 for workers age 16 to 64 (during the earnings year) with 0 to 39 years of potential experience whose class of work in their longest job was private or government wage/salary employment. Hourly earnings are calculated as annual earnings divided by the product of weeks worked and usual hours in the prior year. Full-time, full-year workers are those who work 35 hours per week (using the Census Bureau's full-time worker flag) and worked 40-plus weeks in the previous year. Full-time weekly earnings are calculated as the logarithm of annual earnings over weeks worked for the full-time, full-year sample. Allocated earnings observations are excluded after 1966 using family earnings allocation flags (1967 to 1974) or individual earnings allocation flags (1975 forward). Weights are used in all calculations. Full-time earnings are weighted by the product of the CPS sampling weight and weeks worked. Hourly earnings are weighted by the product of the CPS sampling weight, weeks worked, and hours worked in the prior year.

Prior to March 1989, all wage and salary income in the March CPS was reported in a single variable, which was topcoded at values between $50,000 and $99,999 in years 1964 to 1988. For these cases, we multiply the topcoded earnings value by 1.5, following Katz and Murphy (1992). Commencing in 1989, wage and salary incomes were collected in two separate earnings variables, corresponding to primary and secondary labor earnings. After adjusting for topcoding, we sum these values to calculate total wage and salary earnings. Topcodes after 1988 are handled as follows. For the primary earnings variable, topcoded values are reported at the topcode maximum up to 1996. We multiply these values by 1.5. Starting in 1996, topcoded primary earnings values are assigned the mean of all topcoded earners. In these cases, we simply reassign the topcoded value and, again, multiply by 1.5. For the secondary earnings value, the topcoded

maximum is set at 99,999 from 1989 to 1995 and then falls to 25,000 in 1996 forward. For lack of a superior alternative, we again use the topcoded value multiplied by 1.5.

After making adjustments for topcoding, full-time earnings of below $67/week in 1982$ ($112/week in 2000$) and hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000$) are dropped, as are hourly wages exceeding 1/35th the topcoded value of weekly earnings.

C. Coding of education and potential experience in CPS samples

To attain comparable educational categories across the redefinition of Census Bureau's education variable introduced in 1992 in the CPS, we use the method proposed by Jaeger (1997). In CPS samples coded with the pre-1992 education question, we defined high school dropouts as those with fewer than 12 years of completed schooling; high school graduates as those having 12 years of completed schooling; some college attendees as those with any schooling beyond 12 years (completed or not) and less than 16 completed years; and college plus graduates as those with 16 or more years of completed schooling. In CPS samples coded with the revised education question, we define high school dropouts as those with fewer than 12 years of completed schooling; high school graduates as those with either 12 completed years of schooling and/or a high school diploma or G.E.D.; some college as those attending some college or holding an Associate's Degree; and college plus as those with a B.A. or higher.

To calculate potential experience in data years coded with the revised education question, we use figures from Park (1994) to assign years of completed education to each worker based upon race, gender and highest degree held. Years of potential experience were calculated as age minus assigned years of education minus 6, rounded down to the nearest integer value.

D. Construction of relative wage series

We calculate composition-adjusted college-high school relative wages overall and by age or experience using the March and May/ORG samples described above. These data were sorted into sex-education-experience groups based on a breakdown of the data into 2 sexes, 5 education categories (high school dropout, high school graduate, some college, college plus, and greater than college), and 4 potential experience categories (0-9, 10-19, 20-29, and 30+ years). Log weekly wages of full-time, full-year workers (March CPS) and all hourly workers (May/ORG) were regressed in each year separately by sex on the dummy variables for 4 education categories, a quartic in experience, 3 region dummies, black and other race dummies, and interactions of the experience quartic with 3 broad education categories (high school graduate, some college, and college plus). The (composition-adjusted) mean log wage for each of the 40 groups in a given year is the predicted log wage from these regressions evaluated for whites, living in the mean geographic region, at the relevant experience level (5, 15, 25 or 35 years depending on the experience group). Mean log wages for broader groups in each year represent weighted averages of the relevant (composition-adjusted) cell means using a fixed set of weights, equal to the mean share of total hours worked by each group over 1967 to 2002 from the March CPS.

E. Construction of relative supply measures

We calculate college/high-school relative supply measures using the March and May/ORG samples described above. We first form a labor 'quantity sample' equal to total hours of worked by all employed workers (including those in self-employment) with 0 to 39 years of potential experience in 400 gender × education × potential experience cells: experience groups are single-year categories of 0 to 39 years; education groups are high school dropout, high school graduate, some college, college graduate, and post-college. The quantity data are merged to a corresponding 'price sample' containing real mean full-time weekly (March CPS) or real mean hourly (May/ORG CPS) earnings by year, gender, potential experience

and education. (Wage data used for the price sample correspond to earnings samples described above.) We normalize wages in each of the 400 earnings cells in each year to an 'efficiency units' measure by dividing by the wage of high-school graduate males with 10 years of potential experience in the contemporaneous year. This normalization yields a relative wage measure for each earnings group in each year; the choice of the base earnings group is innocuous.

The quantity and price samples are combined to calculate relative log college/high-school supplies. Define the efficiency units of labor supply of a gender × education × potential experience group in year t as the efficiency unit wage measure for that group multiplied by the group's quantity of labor supply in year t. Following Autor, Katz and Krueger (1998) and Card and Lemieux (2001), we calculate aggregate college-equivalent labor supply as the total efficiency units of labor supplied by college or college-plus workers plus half of the efficiency units of labor supplied by workers with some college. Similarly, aggregate high-school equivalent labor supply is the sum of efficiency units supplied by high-school or lower workers, plus half of the efficiency units supplied by workers with some college. Our college/high-school log relative supply index is the natural logarithm of the ratio of college-equivalent to non-college equivalent labor supply in each year. This measure is calculated overall for each year and by 10 year potential experience groupings. For relative supply calculations using age instead of potential experience (Appendix Table 2), we repeat this procedure, replacing the 40 potential experience categories by 40 age groups: 25 to 64.

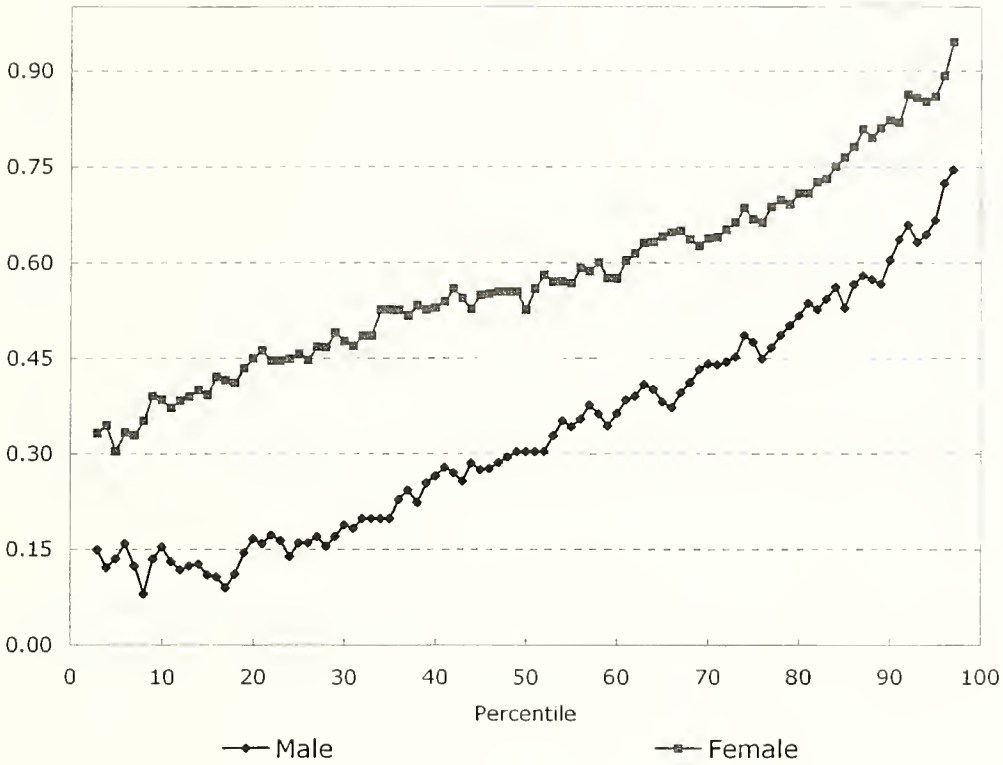


Figure 1. Change in Log Real Weekly Wage by Percentile, Full Time Workers, 1963 - 2003 (March CPS)

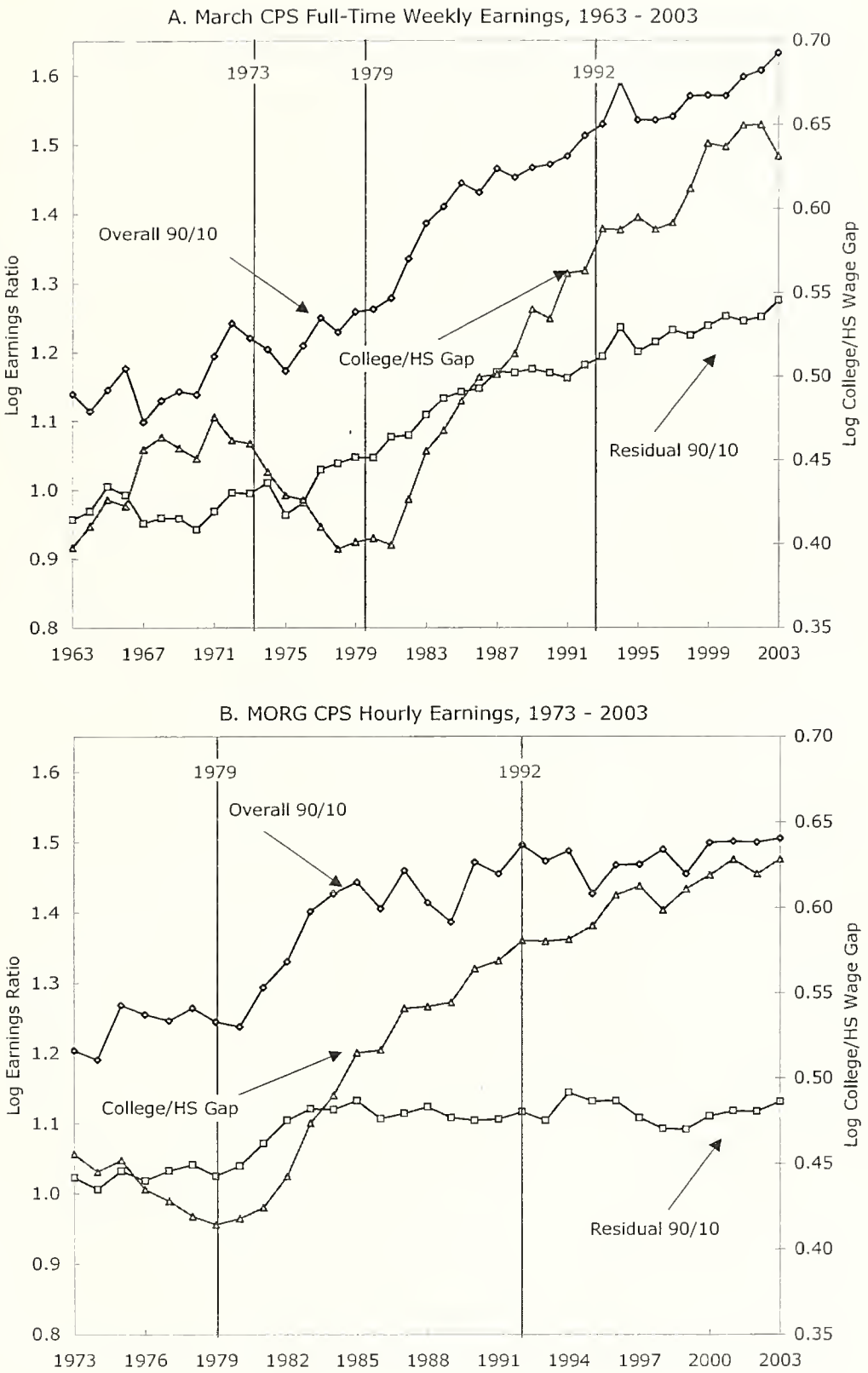


Figure 2. Three Measures of Wage Inequality: College/High School Premium, Male 90/10 Overall Inequality and Male 90/10 Residual Inequality

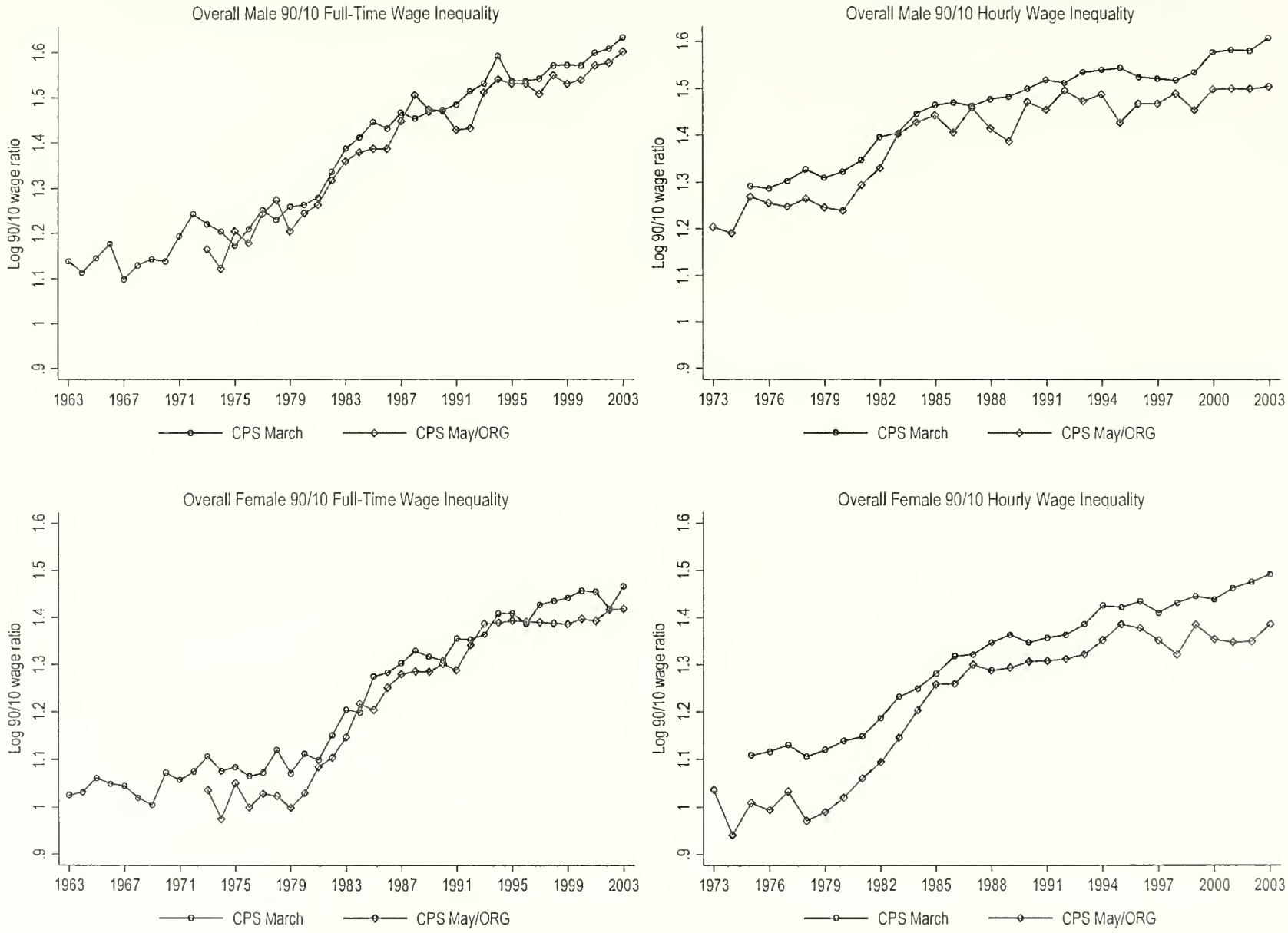


Figure 3. 90/10 Weekly and Hourly Wage Inequality in May/ORG and March CPS Series, 1963 - 2003

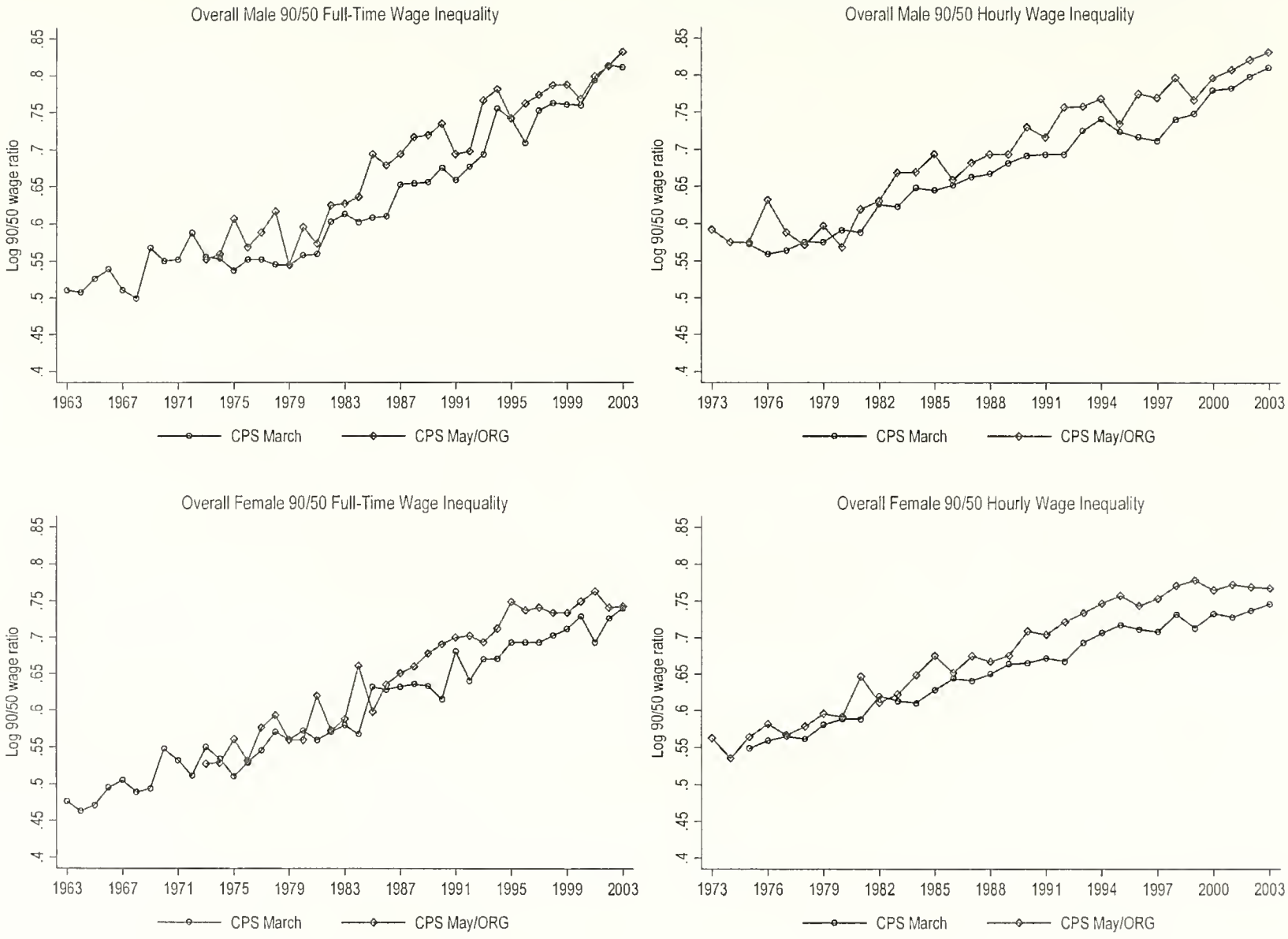


Figure 4. 90/50 Full Time Weekly and Hourly Wage Inequality in May/ORG and March CPS Series, 1963 - 2003

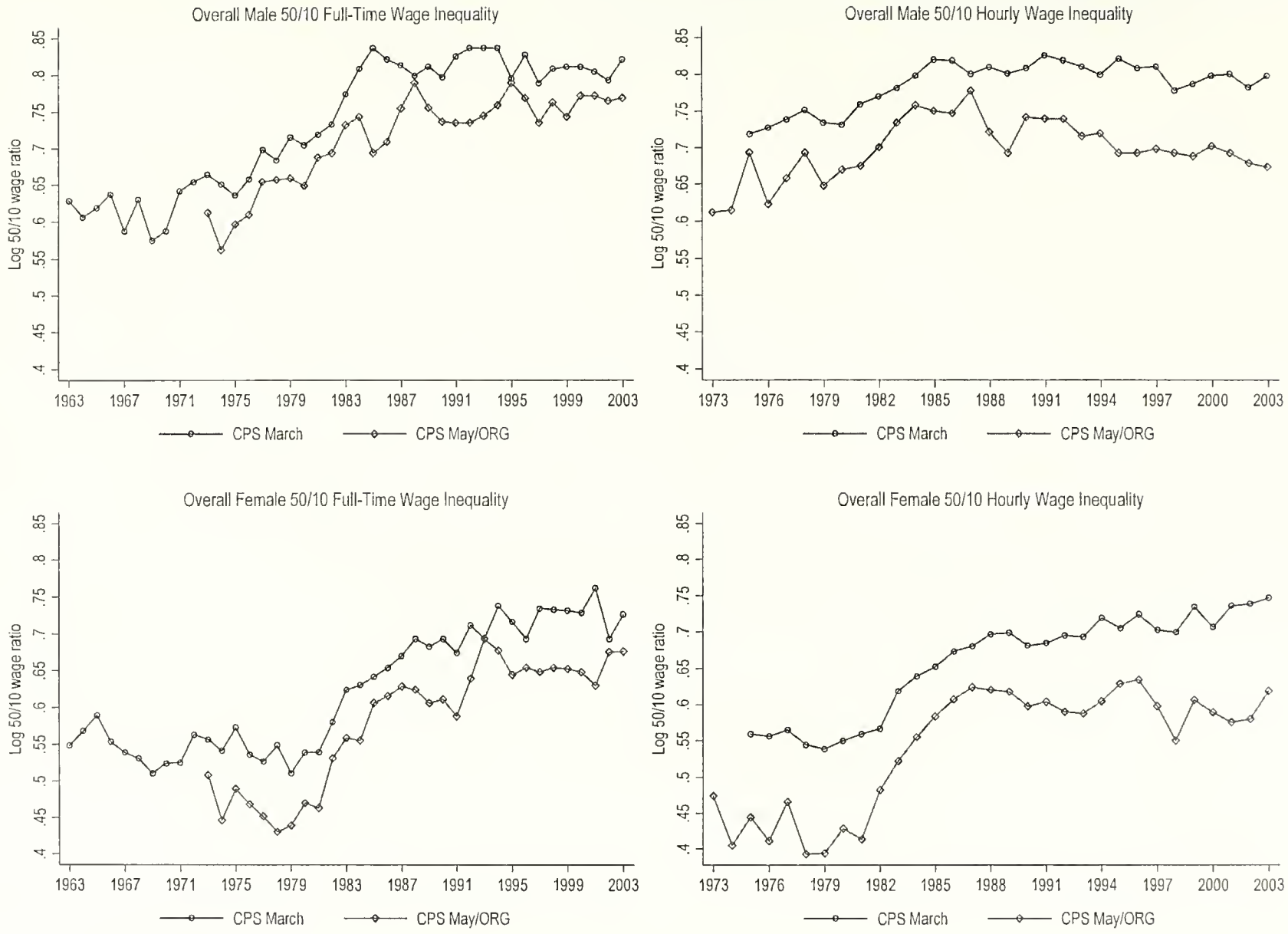


Figure 5. 50/10 Full-Time Weekly and Hourly Wage Inequality in May/ORG and March CPS Series, 1963 - 2003

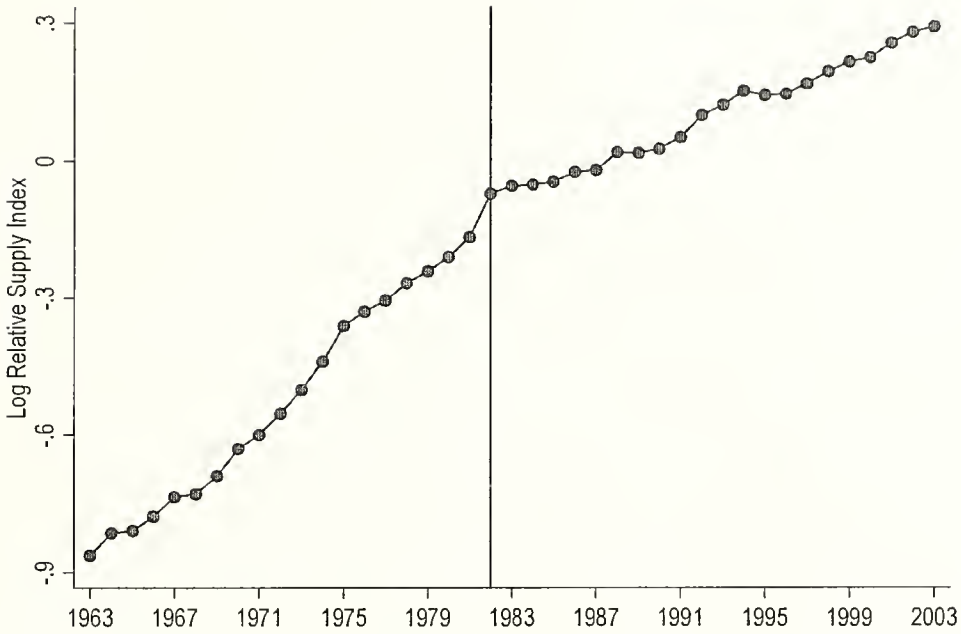
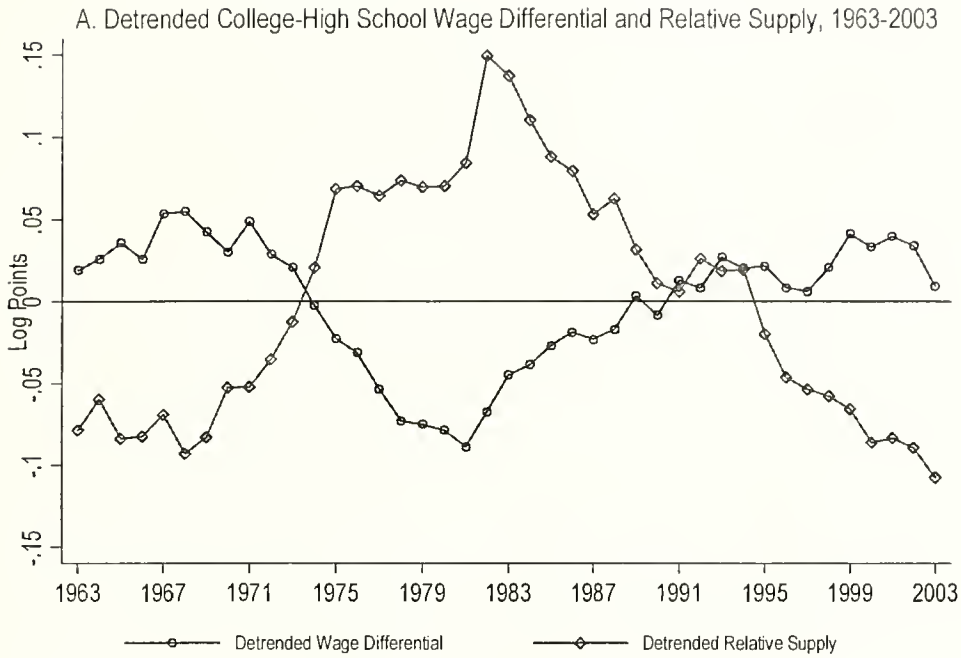


Figure 6. Relative Supply of College Equivalent Labor
1963 - 2003 (March CPS)



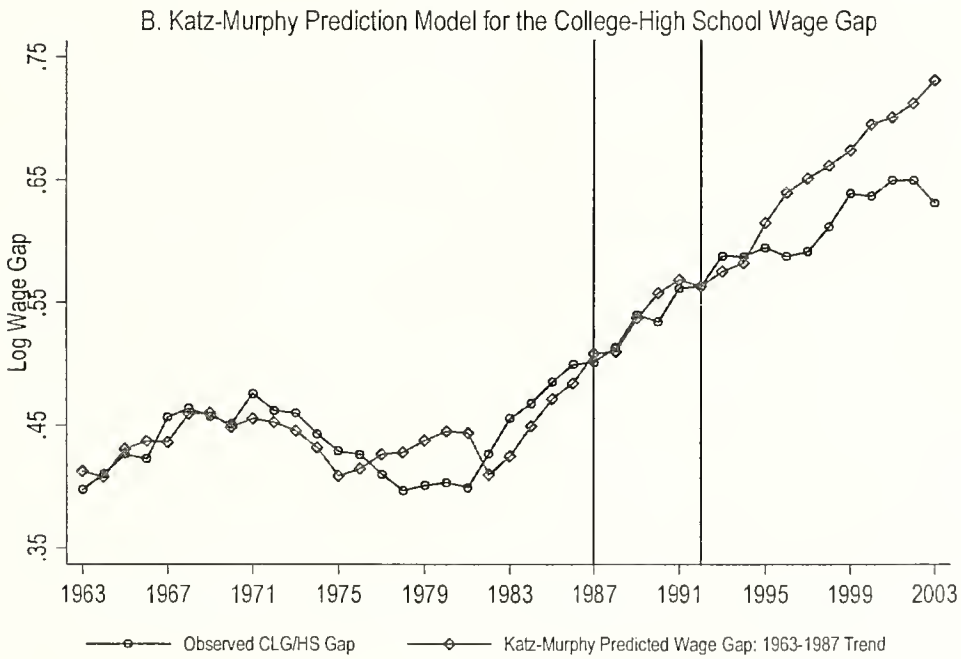
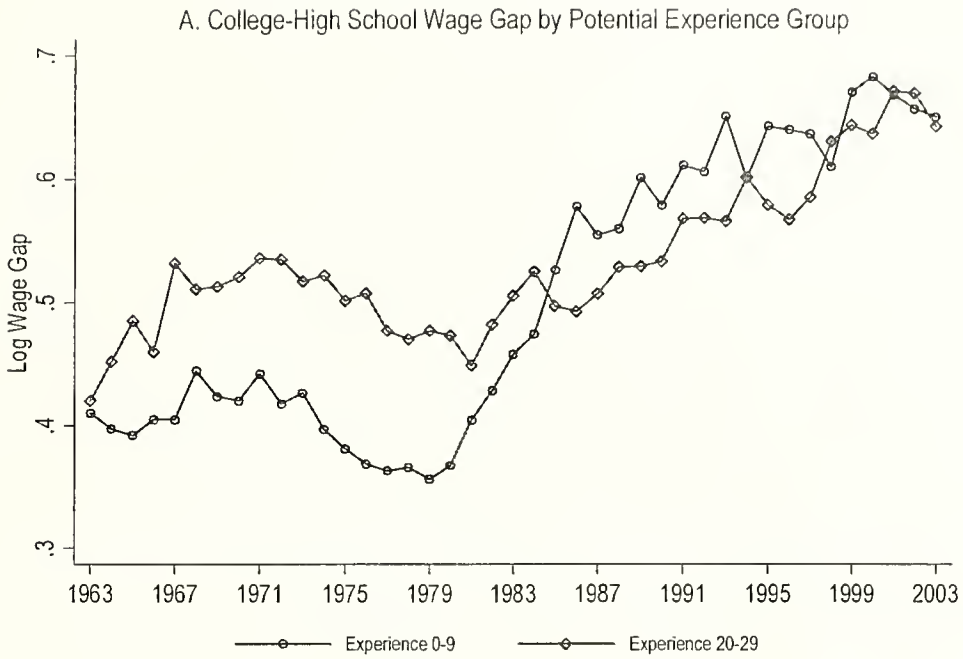


Figure 7. College/High School Relative Supply and Wage Differential, 1963 - 2003 (March CPS) — Panel A. Detrended Measures of Supply and Wage Differentials — Panel B. Katz-Murphy Model for College/HS Wage Differential: Observed and Predicted Values Using 1963 - 1987 Model Estimation



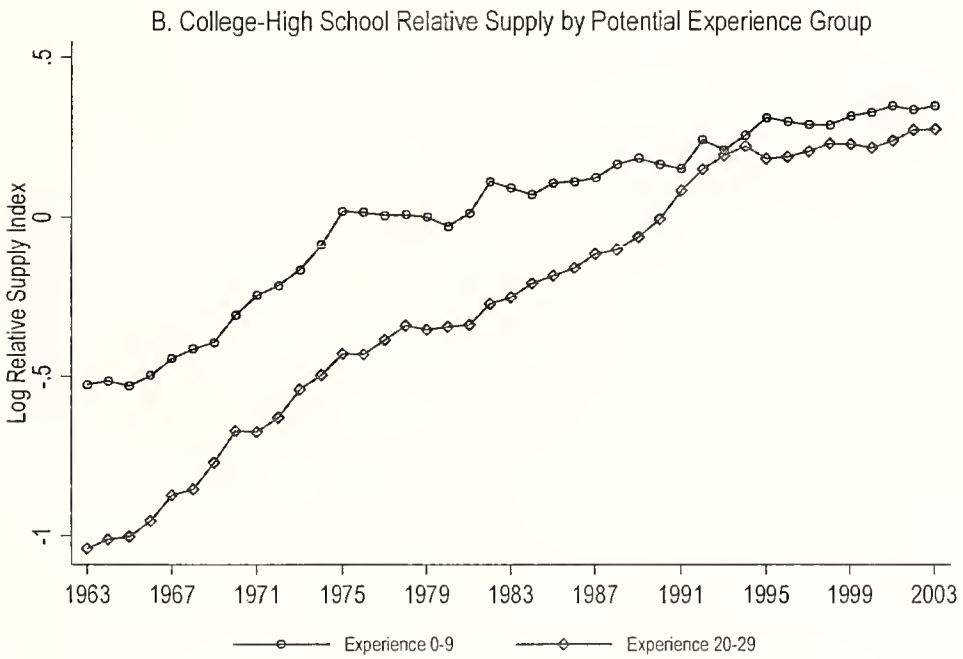


Figure 8. Composition Adjusted Log Relative College/High Wage and Supply by Potential Experience and Age Groups, 1963 - 2003 (March CPS)

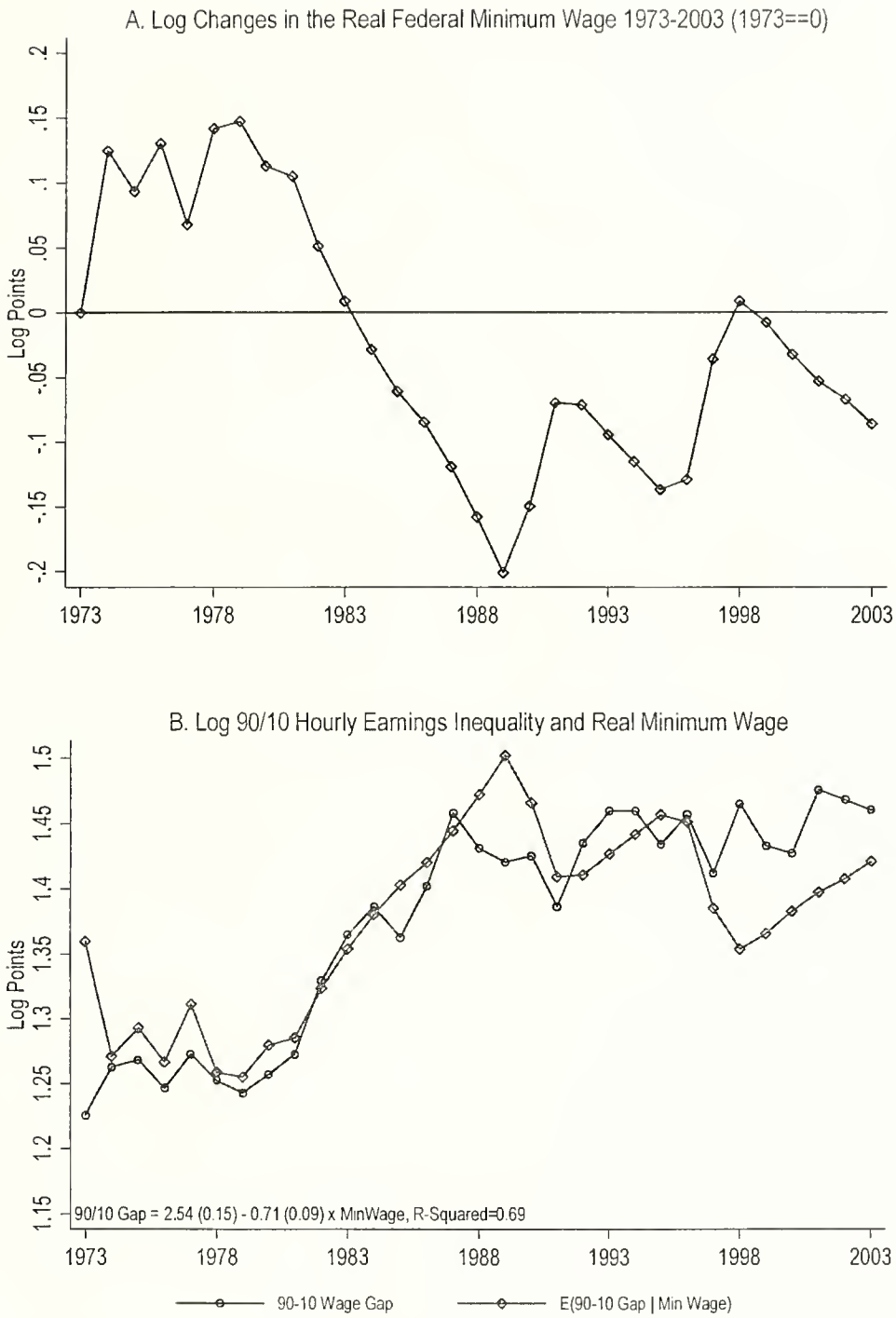
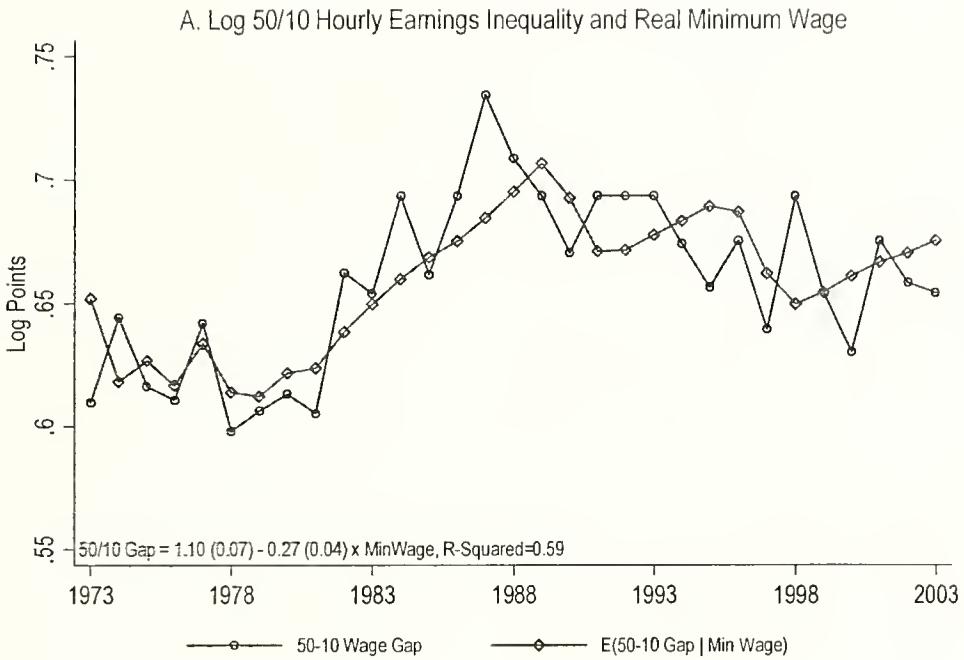


Figure 9. Log Real Federal Minimum Wage and Log 90/10 Hourly Wage Differential, 1973-2003 (May/ORG CPS)



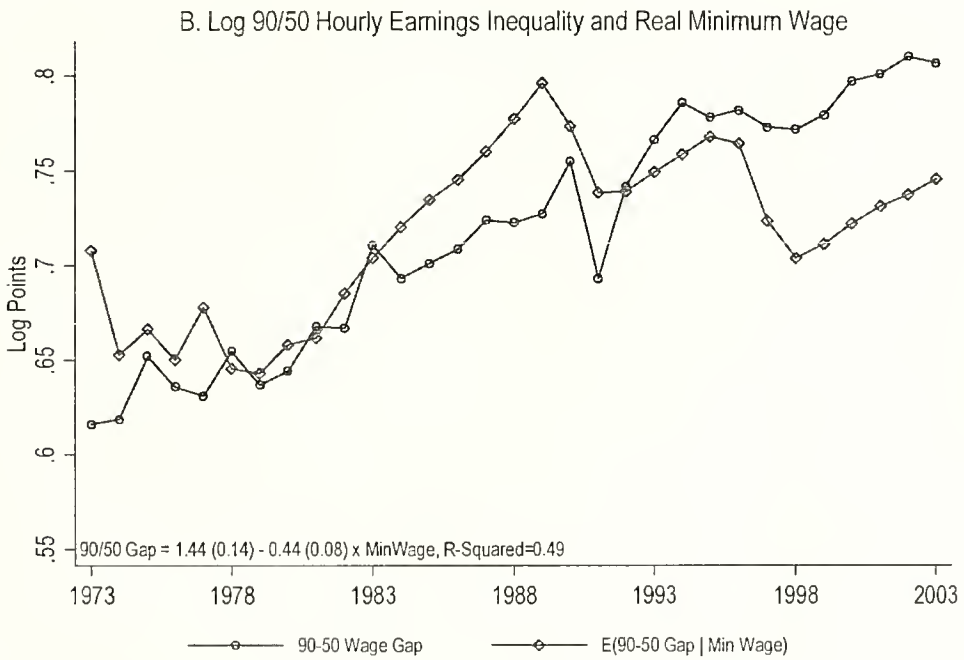


Figure 10. Log 50/10 and 90/50 Hourly Wage Differentials and Log Real Federal Minimum Wage, 1973-2003 (May/ORG CPS)

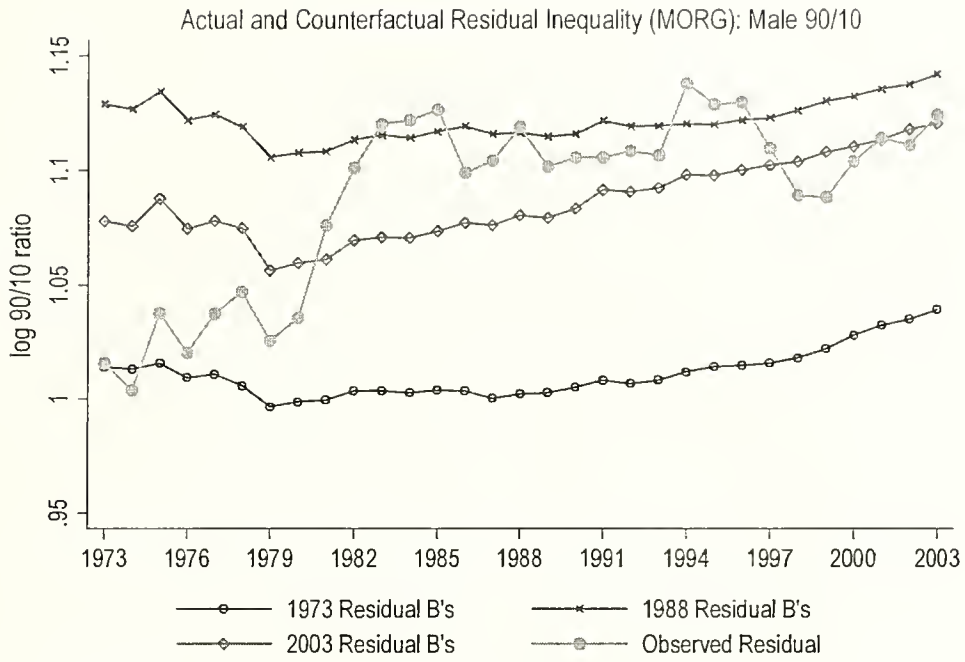


Figure 11a. Actual and Counterfactual Male 90/10 Residual Hourly Wage Inequality 1973-2003 (CPS May/ORG)

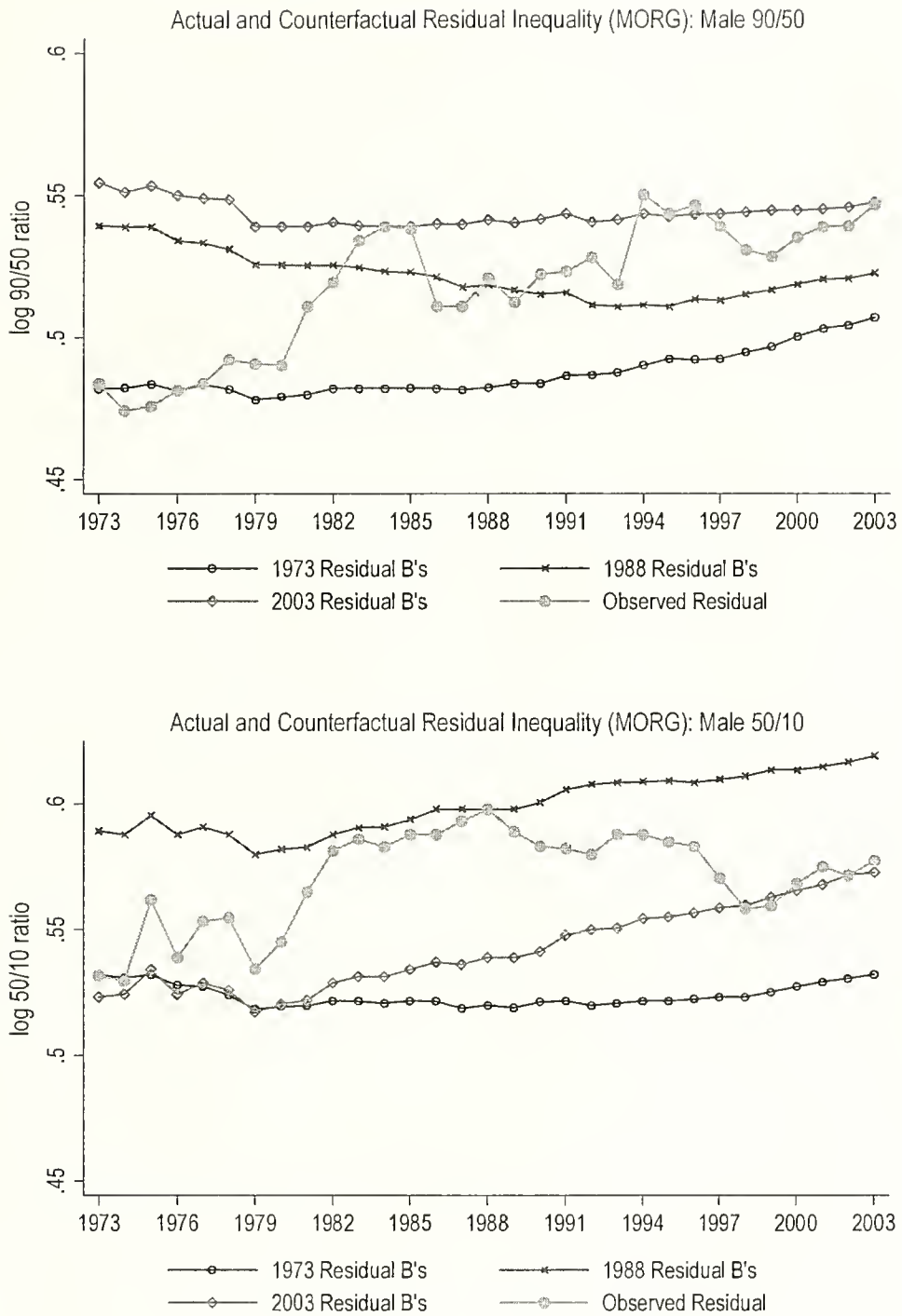


Figure 11b. Actual and Counterfactual Male 90/50 and 50/10 Residual Hourly Wage Inequality 1973-2003 (CPS May/ORG)

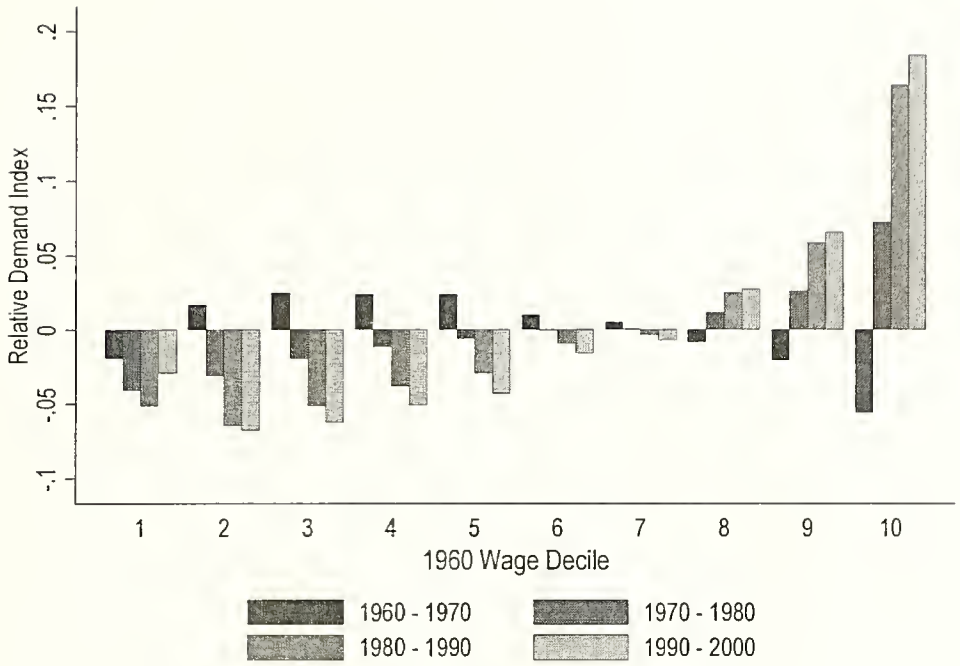


Figure 12. Estimated Relative Employment Demand Shifts by Wage Decile over Four Decades: 1960 - 2000

Table 1. Changes in Real, Composition-Adjusted Log Weekly wages for Full-Time, Full-Year Workers, 1963 - 2003.
(100 x Change in Mean Log Real Weekly Wages)

	1963-1971	1971-1979	1979-1987	1987-1995	1995-2003	1963-1987	1963-2003
All	19.5	0.7	-0.9	-4.9	9.6	14.3	24.0
Sex							
Men	21.1	0.1	-5.0	-7.8	8.1	8.4	16.5
Women	17.2	1.4	4.8	-0.8	11.8	22.6	34.4
Education (years of schooling)							
0-11	17.0	1.8	-8.4	-10.3	5.3	0.1	5.4
12	17.6	3.2	-3.3	-6.6	9.3	11.0	20.2
13-15	18.6	0.5	1.1	-5.3	10.5	14.9	25.3
16+	25.4	-4.3	6.8	2.7	12.9	30.7	43.6
16-17	22.8	-4.9	5.6	1.0	12.8	24.5	37.3
18+	31.4	-2.7	9.5	6.8	13.1	44.9	58.0
Experience (Males)							
5 years	20.0	-3.6	-8.5	-7.6	11.8	0.2	12.0
25-35 years	21.5	3.4	-1.7	-8.2	4.7	15.1	19.8
Education and Experience (Males)							
Education 12							
Experience 5	19.4	0.7	-16.1	-10.3	10.9	-6.4	4.5
Experience 25 - 35	17.0	6.3	-2.5	-7.6	3.7	13.2	16.9
Education 16+							
Experience 5	23.1	-11.0	9.3	-1.9	11.5	19.5	31.0
Experience 25 - 35	35.0	1.7	2.6	-2.2	10.7	37.1	47.8

Tabulated numbers are changes in the (composition-adjusted) mean log wage for each group, using data on full-time, full-year workers ages 16 to 64 from the March CPS covering earnings in calendar years 1963 to 2003. The data were sorted into sex-education-experience groups based on a breakdown of the data into 2 sexes, 5 education categories (high school dropout, high school graduate, some college, college graduate, and post-college), and 4 potential experience categories (0-9, 10-19, 20-29, and 30-39 years). Log weekly wages of full-time, full-year workers were regressed in each year separately by sex on the dummy variables for 4 education categories, a quartic in experience, 3 region dummies, black and other race dummies, and interactions of the experience quartic with 3 broad education categories (high school graduate, some college, and college plus). The (composition-adjusted) mean log wage for each of the 40 groups in a given year is the predicted log wage from these regressions evaluated for whites, living in the mean geographic region, at the relevant experience level (5, 15, 25 or 35 years depending on the experience group). Mean log wages for broader groups in each year represent weighted averages of the relevant (composition-adjusted) cell means using a fixed set of weights, equal to the mean share of total hours worked by each group over 1963 - 2003. All earnings numbers are deflated by the chain-weighted (implicit) price deflator for personal consumption expenditures. Earnings of less than $67/week in 1982$ ($112/week in 2000$). Allocated earnings observations are excluded in years 1967 forward using either family earnings allocation flags (1967 - 1974) or individual earnings allocation flags (1975 forward).

Table 2. Regression Models for the College/High School Log Wage Gap, 1963 - 2003

	(1)	(2)	(3)	(4)	(5)	(6)
	1963 - 1987	1963- 2003				
CLG/HS relative supply	-0.638 (0.129)	-0.462 (0.047)	-0.612 (0.065)	-0.636 (0.095)	-0.498 (0.069)	
Log real minimum wage				-0.057 (0.050)	-0.113 (0.043)	-0.146 (0.066)
Male prime-age unemp. rate				0.003 (0.003)	0.002 (0.004)	-0.018 (0.003)
Time	0.026 (0.005)	0.020 (0.001)	0.025 (0.002)	0.025 (0.003)	0.020 (0.002)	0.006 (0.001)
Time x Post-1992			-0.007 (0.002)	-0.005 (0.003)		
Constant	-0.165 (0.120)	-0.004 (0.039)	-0.142 (0.057)	-0.065 (0.154)	0.166 (0.108)	0.694 (0.123)
Observations	25	41	41	41	41	41
R-squared	0.571	0.936	0.949	0.953	0.948	0.873

Standard errors in parentheses. Each column presents an OLS regression of the fixed-weighted college high school wage differential on the indicated variables. Real minimum wage is deflated by the Personal Consumption Expenditure Deflator. Source data for supply and earnings measures: March CPS 1964 - 2004.

Table 3. Regression Models for the College/High School Log Wage Gap by Potential Experience Group, 1963-2003, Males and Females Pooled.

	Potential Experience Groups					
	All Experience Groups		0-9 yrs	10-19 yrs	20-29 yrs	30-39 yrs
Own Supply Minus Aggregate Supply	-0.292	-0.293	-0.106	-0.257	0.217	-0.030
	(0.028)	(0.028)	(0.098)	(0.074)	(0.088)	(0.108)
Aggregate Supply	-0.613	-0.655	-0.837	-0.684	-0.151	-0.396
	(0.055)	(0.081)	(0.129)	(0.103)	(0.142)	(0.164)
Log Real Minimum Wage		-0.028	-0.237	-0.145	0.026	0.103
		(0.042)	(0.068)	(0.057)	(0.068)	(0.082)
Prime Age Male Unemployment		0.003	0.007	0.006	0.000	-0.002
		(0.003)	(0.005)	(0.004)	(0.004)	(0.006)
Time	0.025	0.026	0.033	0.025	0.006	0.017
	(0.002)	(0.003)	(0.005)	(0.004)	(0.005)	(0.006)
Time x Post-1992	-0.005	-0.005	-0.012	0.002	0.007	-0.008
	(0.002)	(0.002)	(0.003)	(0.003)	(0.004)	(0.005)
Constant	-0.044	-0.037	0.097	0.138	0.335	-0.062
	(0.046)	(0.129)	(0.207)	(0.175)	(0.229)	(0.262)
N	164	164	41	41	41	41
R-squared	0.865	0.867	0.955	0.968	0.876	0.645

Standard errors in parentheses. Each column presents an OLS regression of the fixed-weighted college/high school wage differential on the indicated variables. The college/high school wage premium is calculated at the mid-point of each potential experience group. See Notes to Table 1 for information on fixed-weighting scheme. Real minimum wage is deflated by the Personal Consumption Expenditure Deflator. Columns 1 and 2 also include dummy variables for the 4 potential experience groups used in the table.

Table 4. Regression Models for Log Earnings Gaps: Males and Females combined, 1973 - 2003 (May/ORG CPS)

	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
	90/10 Earnings Gap				90/50 Earnings Gap				50/10 Earnings Gap			
CLG/HS relative supply	0.365		0.186	0.160	0.070		0.028	0.054	0.295		0.158	0.107
	(0.154)		(0.140)	(0.149)	(0.068)		(0.089)	(0.094)	(0.132)		(0.138)	(0.143)
Log real minimum wage		-0.348	-0.343	-0.315		-0.084	-0.083	-0.109		-0.265	-0.260	-0.206
		(0.056)	(0.055)	(0.073)		(0.034)	(0.035)	(0.046)		(0.054)	(0.054)	(0.071)
Male prime-age unemp. rate		0.007	0.003	0.003		0.001	0.001	0.001		0.005	0.002	0.002
		(0.003)	(0.004)	(0.004)		(0.002)	(0.003)	(0.003)		(0.003)	(0.004)	(0.004)
Time	0.000	0.006	0.002	0.003	0.005	0.006	0.005	0.004	-0.005	0.000	-0.004	-0.001
	(0.004)	(0.001)	(0.003)	(0.004)	(0.002)	(0.000)	(0.002)	(0.003)	(0.003)	(0.001)	(0.003)	(0.004)
Time x Post-1992				-0.002				0.002				-0.004
				(0.003)				(0.002)				(0.003)
Constant	1.377	1.820	1.899	1.838	0.637	0.752	0.764	0.824	0.740	1.068	1.135	1.014
	(0.059)	(0.103)	(0.117)	(0.159)	(0.026)	(0.063)	(0.075)	(0.101)	(0.050)	(0.100)	(0.115)	(0.153)
Observations	31	31	31	31	31	31	31	31	31	31	31	31
R-squared	0.843	0.934	0.938	0.939	0.943	0.953	0.953	0.955	0.328	0.632	0.650	0.668

Standard errors in parentheses. Each column presents a separate OLS regression using May/ORG CPS data. The real minimum wage series is deflated by the Personal Consumption Expenditure Deflator.

Appendix Table 1a. Trends in Overall Inequality 1975 to 2003.
100 x Changes in Inequality Measures

A. 90th Percentile - 10th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	10.5	5.7	5.6	2.5	1.5	3.4	3.9	5.2	1.9	2.8	4.3	0.4
1981 - 1988	17.4	24.3	13.0	12.1	23.1	20.1	19.9	22.7	14.8	16.7	9.4	15.8
1988 - 1994	13.8	3.5	6.3	7.4	8.0	10.4	7.9	6.4	5.0	2.2	4.3	2.8
1994 - 2003	4.1	6.1	6.9	1.8	5.8	3.0	6.7	3.6	6.9	6.8	5.3	0.1
1975 - 1988	28.0	30.0	18.6	14.6	24.6	23.5	23.9	27.9	16.8	19.5	13.6	16.3
1988 - 2003	17.9	9.6	13.2	9.1	13.7	13.4	14.6	9.9	11.9	8.9	9.5	2.9

B. 90th Percentile - 50th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	2.2	-3.3	1.5	4.3	4.9	5.9	3.9	8.3	2.7	1.9	4.5	1.6
1981 - 1988	9.4	14.2	7.9	7.5	7.6	4.0	6.2	2.0	5.3	8.5	1.4	5.5
1988 - 1994	10.1	6.5	7.4	7.5	3.5	5.2	5.6	8.0	6.0	1.3	4.3	6.3
1994 - 2003	5.6	5.1	7.0	6.3	6.9	3.1	4.0	2.1	6.4	5.8	7.3	2.1
1975 - 1988	11.7	10.9	9.4	11.8	12.5	10.0	10.1	10.3	8.0	10.4	5.9	7.0
1988 - 2003	15.7	11.6	14.4	13.8	10.4	8.2	9.6	10.1	12.4	7.1	11.6	8.4

C. 50th Percentile - 10th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	8.3	9.0	4.1	-1.8	-3.4	-2.6	0.0	-3.1	-0.8	0.9	-0.3	-1.1
1981 - 1988	8.0	10.1	5.1	4.6	15.4	16.1	13.7	20.7	9.5	8.2	8.0	10.3
1988 - 1994	3.8	-3.0	-1.1	-0.2	4.4	5.2	2.2	-1.6	-1.0	0.8	0.0	-3.5
1994 - 2003	-1.5	1.0	-0.1	-4.5	-1.2	-0.1	2.8	1.5	0.5	1.0	-2.0	-2.0
1975 - 1988	16.3	19.1	9.2	2.8	12.0	13.5	13.7	17.6	8.8	9.1	7.7	9.2
1988 - 2003	2.2	-2.0	-1.2	-4.7	3.3	5.2	5.0	-0.2	-0.6	1.9	-2.0	-5.4

D. Variance

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	3.4	2.8	2.8	1.3	1.3	0.9	0.7	0.9	1.8	1.1	1.2	0.6
1981 - 1988	7.4	6.8	5.1	4.9	6.8	7.1	6.5	6.4	5.0	5.3	3.9	4.4
1988 - 1994	4.7	3.6	3.5	2.5	5.0	3.8	4.5	3.7	3.7	2.9	3.1	2.2
1994 - 2003	3.1	0.6	2.6	-0.3	3.2	1.6	3.2	0.9	2.4	0.5	2.3	0.0
1975 - 1988	10.8	9.5	7.9	6.3	8.0	8.0	7.2	7.3	6.8	6.5	5.1	5.0
1988 - 2003	7.8	4.3	6.1	2.3	8.2	5.5	7.7	4.6	6.2	3.4	5.4	2.2

Appendix Table 1b. Trends in Residual Inequality 1975 to 2003
100 x Changes in Inequality Measures

A. 90th Percentile - 10th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	11.3	9.0	8.0	3.6	5.2	5.9	2.4	2.6	8.0	6.4	5.4	3.1
1981 - 1988	9.3	9.2	4.9	4.3	13.2	13.3	12.8	13.1	11.3	10.4	7.7	7.4
1988 - 1994	6.5	5.9	4.9	2.3	5.1	4.8	4.3	2.7	4.4	5.5	4.0	2.4
1994 - 2003	3.9	2.8	3.4	-1.1	3.3	0.9	1.9	0.3	3.0	1.1	2.2	-1.6
1975 - 1988	20.7	18.2	12.9	7.9	18.3	19.2	15.2	15.7	19.3	16.8	13.0	10.5
1988 - 2003	10.4	8.7	8.3	1.2	8.4	5.7	6.1	3.0	7.4	6.6	6.2	0.8

B. 90th Percentile - 50th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	5.6	5.1	3.8	3.5	2.6	6.8	1.8	4.2	4.1	3.8	3.5	4.3
1981 - 1988	3.6	2.9	1.9	0.2	5.4	1.5	3.6	1.7	4.5	3.4	2.8	1.1
1988 - 1994	4.7	5.0	5.3	3.9	2.9	3.5	3.8	3.1	3.6	3.8	4.2	3.2
1994 - 2003	4.1	3.0	2.7	-0.1	2.5	1.8	1.6	-0.8	3.1	2.5	2.0	-0.4
1975 - 1988	9.2	8.0	5.7	3.7	8.0	8.3	5.4	5.9	8.6	7.3	6.4	5.4
1988 - 2003	8.8	8.0	7.9	3.8	5.4	5.3	5.4	2.3	6.7	6.4	6.2	2.7

C. 50th Percentile - 10th Percentile

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	5.7	3.9	4.2	0.1	2.6	-0.9	0.6	-1.6	3.9	2.6	1.8	-1.2
1981 - 1988	5.8	6.3	3.0	4.1	7.7	11.7	9.2	11.4	6.8	6.9	4.8	6.3
1988 - 1994	1.8	0.9	-0.3	-1.6	2.2	1.3	0.4	-0.4	0.7	1.7	-0.2	-0.8
1994 - 2003	-0.2	-0.2	0.7	-1.1	0.8	-0.9	0.3	1.1	0.0	-1.4	0.2	-1.1
1975 - 1988	11.5	10.3	7.2	4.2	10.4	10.9	9.8	9.8	10.8	9.5	6.7	5.1
1988 - 2003	1.6	0.7	0.4	-2.7	3.0	0.4	0.7	0.7	0.7	0.3	0.0	-1.9

D. Variance

	Males				Females				Overall			
	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly	March FT	MORG FT	March Hourly	MORG Hourly
1975 - 1981	2.9	2.2	2.4	0.5	1.4	1.3	0.9	0.8	2.2	1.7	1.7	0.3
1981 - 1988	3.9	3.1	2.4	1.5	4.4	4.0	3.9	3.1	3.9	3.4	2.9	2.0
1988 - 1994	2.2	2.5	1.2	1.5	2.2	2.1	2.1	2.5	2.1	2.2	1.5	1.7
1994 - 2003	2.2	-0.6	2.0	-1.3	2.1	0.2	1.9	-0.7	1.9	-0.4	1.8	-1.1
1975 - 1988	6.7	5.4	4.8	2.0	5.8	5.3	4.8	3.9	6.1	5.0	4.5	2.3
1988 - 2003	4.4	2.0	3.2	0.2	4.2	2.2	4.1	1.8	4.0	1.8	3.2	0.6

Table Notes:

Appendix Tables 1a and 1b:

Data sources for May/ORG statistics are May CPS for 1976 to 1978 and CPS Merged Outgoing Rotation Groups for years 1979 to 2003. Samples include wage/salary workers ages 16 - 64 with 0 - 38 years of potential experience in current employment. Full-time earnings is the logarithm of reported usual weekly earnings. Hourly wages are the logarithm of reported hourly earnings for those paid by the hour and the logarithm of usual weekly earnings divided by hours worked last week for non-hourly workers. Top-coded earnings observations are multiplied by 1.5. Full-time earnings of below $67/week in 1982$ ($112/week in 2000$) and hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000$) are dropped, as are hourly wages exceeding 1/35th the top-coded value of weekly earnings. Allocated earnings observations are excluded in all years, except where allocation flags are unavailable (1994 and 1995). Full-time earnings are weighted by CPS sampling weights. Hourly earnings are weighted by the product of CPS sampling weights and hours worked in the prior week.

Data sources for March series is the March Current Population for earnings years 1975 to 2003 for workers age 16 - 64 (during earnings year) with 0 - 38 years of potential experience whose class of work in their longest job was private or government wage/salary employment. Hourly earnings are calculated as annual earnings divided by the product of weeks worked and usual hours in the prior year. Full-time, full-year workers are those who work 35 hours per week and worked 40+ weeks in the previous year. Full-time weekly earnings are calculated as the logarithm of annual earnings over weeks worked for the full-time, full-year sample. Top-coded earnings observations are multiplied by 1.5. Full-time earnings of below $67/week in 1982$ ($112/week in 2000$) and hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000$) are dropped, as are hourly wages exceeding 1/35th the top-coded value of weekly earnings. Allocated earnings observations are excluded after 1966 using family earnings allocation flags (1967 - 1974) or individual earnings allocation flags (1975 forward).

Appendix Table 1a:

Tabulated numbers are 100 times changes in aggregate inequality statistics by gender and overall for full-time and all hourly workers.

Appendix Table 1b:

Tabulated numbers are 100 times changes in indicated residual inequality statistics by gender and overall for full-time and all hourly workers. Wage residuals are calculated from a regression of the indicated weekly or hourly wage measure on 5 education category dummies interacted with 13 dummies categorizing potential experience (64 dummies and 1 omitted). Education categories are high school dropout, high school graduate, some college, exactly college graduate, and post-college. Potential experience categories are 0-2 years, 3-5 years ... 36-38 years. Pooled-gender regressions also include a female dummy.

Appendix Table 2. Regression Models for the College/High School Log Wage Gap by Age Group, 1963-2003, Males and Females Pooled.

	Age Groups					
	All Age Groups		Age 25-34	Age 35-44	Age 45-54	Age 55-64
Own Supply Minus Aggregate Supply	-0.209	-0.209	-0.159	-0.102	-0.165	-0.044
	(0.017)	(0.017)	(0.053)	(0.102)	(0.106)	(0.115)
Aggregate Supply	-0.511	-0.549	-0.659	-0.519	-0.565	-0.306
	(0.037)	(0.054)	(0.091)	(0.099)	(0.146)	(0.179)
Log Real Minimum Wage		0.020	-0.065	-0.015	0.192	0.027
		(0.034)	(0.049)	(0.075)	(0.068)	(0.094)
Prime Age Male Unemployment		0.002	0.003	0.005	0.003	-0.005
		(0.002)	(0.003)	(0.004)	(0.005)	(0.007)
Time	0.023	0.024	0.028	0.022	0.026	0.016
	(0.001)	(0.002)	(0.004)	(0.004)	(0.006)	(0.007)
Time x Post-1992	-0.004	-0.005	-0.004	-0.001	-0.009	-0.006
	(0.001)	(0.002)	(0.003)	(0.004)	(0.004)	(0.005)
Constant	-0.157	-0.236	-0.201	-0.028	-0.429	0.130
	(0.035)	(0.100)	(0.127)	(0.194)	(0.244)	(0.275)
N	164	164	41	41	41	41
R-squared	0.932	0.933	0.981	0.926	0.783	0.709

Standard errors in parentheses. Each column presents an OLS regression of the fixed-weighted college/high school wage differential on the indicated variables. The college/high school wage premium is calculated at the mid-point of each age range. See Notes to Table 1 for information on fixed-weighting scheme. Real minimum wage is deflated by the Personal Consumption Expenditure Deflator. Columns 1 and 2 also include dummy variables for the 4 age groups used in the table.

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